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A Returns-Based Representation of Earnings Quality

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ABSTRACT: We examine the properties of a returns-based representation of earnings quality, estimated from firm-specific asset-pricing regressions augmented by an earnings quality mimicking factor. The coefficient on the earnings quality factor (the "e-loading") captures the sensitivity of the firm's returns to earnings quality in a given year or quarter, analogous to beta as a measure of the sensitivity of returns to market movements. Relative to other proxies for earnings quality, e-loadings can be calculated for larger samples of firms and can be estimated for shorter intervals at any point in time. Along all dimensions examined, we find that e-loadings perform well in capturing notions of earnings quality.

Keywords: *earnings quality; returns-based measures; mimicking factor.*

Data Availability: *The earnings quality mimicking factor developed in this study is available upon request. All data used to construct this factor are available from commercial providers (Compustat and CRSP).*

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I. INTRODUCTION

We analyze a returns-based measure of earnings quality that can be estimated for a given firm-year (in extensions, we show the results are generalizable to a firm-quarter). Our measure is the slope coefficient from a regression of a firm's daily excess returns in year *T* on a factor-mimicking portfolio capturing earnings quality, controlling for other factors known to affect returns (market risk premium, size, and book-to-market ratio). We build on the factor-mimicking portfolio approach introduced by Fama and French (1993) and applied by Francis et al. (2005) to demonstrate a risk premium for firms with poorer earnings quality, as captured by Dechow and Dichev's (2002) measure of accruals quality (*AQ*). Specifically, we follow Francis et al.'s (2005) procedures to create an *AQ* factor-mimicking portfolio (*AQfactor*), except that we use daily returns rather than monthly returns to estimate the asset-pricing regressions.[1] Next, for each firm with at least 100 daily returns observations in year *T* = 1970–2003, we estimate one-factor and three-factor annual regressions that add *AQfactor* as an independent variable. Just as the CAPM beta captures exposure to market risk, so too will the coefficient on *AQfactor* in these regressions—the e-loading—capture investor perceptions of the firm's earnings quality exposure in year *T* (a larger e-loading implies greater sensitivity to poor earnings quality).

Capturing the returns consequences of a firm characteristic by applying a factor-mimicking portfolio is not new, nor are construct validity tests of those returns consequences. For example, Fama and French (1993) propose and validate the use of factor-mimicking portfolios and associated loadings as proxies for firm size and book-to-market;[2] and Lamont et al. (2001) probe the existence of a factor capturing the degree to which firms are financially constrained. Our paper extends this body of work by proposing and validating e-loadings as a returns-based representation of earnings quality.

We view earnings quality as a measure of information risk, and we define earnings quality in terms of precision, namely, the mapping of current accruals into current, last year, and next year cash flows. Following Dechow and Dichev (2002), we term this mapping "accruals quality," and denote it *AQ*. Theoretical support for the view that information uncertainty is a nondiversifiable (i.e., priced) risk factor is provided by Easley and O'Hara (2004)[3] and Leuz and Verrecchia (2005). Empirical support for the view that earnings quality as measured by accruals quality is priced is provided by Francis et al. (2005), who show that the returns representation of the *AQ* measure, *AQfactor*, enters one- and three-factor asset-pricing regressions with a reliably positive average coefficient estimate. Thus, conditioning on theory that shows information uncertainty is a priced risk factor, and on

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[1] Using monthly returns in asset-pricing tests results in less estimation error than using daily returns. However, if the true asset-pricing parameters vary over time, then monthly data will obscure this pattern, since a longer time period is needed to estimate an asset-pricing regression with monthly data. Most studies require a minimum of 24 to 60 monthly observations to estimate an asset-pricing regression; consequently, the resulting coefficient estimates are effectively assumed constant over this two- to five-year interval.

[2] The degree to which factor-mimicking portfolios proxy for underlying size and book-to-market constructs continues to be debated. In cross-sectional tests, Daniel and Titman (1997) find that additional variation in returns can be explained by firms' size and book-to-market ratios over and above three-factor returns and loadings. However, using time-series tests, Lewellen (1999) documents no incremental information about expected returns in book-to-market ratios, beyond three-factor returns and loadings.

[3] In Easley and O'Hara's (2004) model, the risk premium associated with information uncertainty is a function of private information (which pertains to information asymmetry) and the precision of public and private information. Our accruals quality measure focuses on the precision of current accruals with respect to cash flows; as such, we view it as capturing the precision of public information. In extensions of our main tests, we consider the overlap between e-loadings and two trading-based measures of information asymmetry: bid-ask spreads (which reflect both public and private information) and probability-of-informed (PIN) trading scores (which arguably focus more on private information).

Francis et al.'s (2005) empirical evidence that the market prices information risk as captured by earnings quality measured by *AQ*, we view e-loadings as capturing the sensitivity of stock returns to earnings quality.

As discussed in more detail in Section II, our paper contributes to the earnings quality literature by establishing an earnings quality metric that offers several advantages relative to traditional earnings quality measures (either measures derived from accounting data or measures derived from both returns and accounting data). First, e-loadings can be measured for firms that lack the time-series of accounting data that is typically required for estimating accounting-based measures of earnings quality. Therefore, e-loadings can be measured for much larger (and more representative of the population) samples of firms, increasing sampling power and generalizability. Second, e-loadings can be reliably measured over intervals as short as a quarter, so they can be used to analyze changes in earnings quality associated with specific events. In addition, the returns-based representation approach is flexible, in that other earnings attributes, such as smoothness and persistence, can also be represented by factor loadings. As we discuss in Section II, these advantages permit researchers to examine earnings quality in settings and for samples that are difficult (or not possible) to examine with traditional measures of earnings quality.

Our analysis of the properties of e-loadings as measures of earnings quality has four components. First, we investigate whether e-loadings vary cross-sectionally with characteristics expected to be related to earnings quality and, separately, with other proxies for earnings quality. Second, we examine whether e-loadings are associated with predictable variation in market participant behavior with respect to earnings. These tests focus on whether investors attach lower earnings response coefficients to higher e-loading firms and whether there is greater dispersion and less accuracy in analysts' earnings forecasts for higher e-loading firms. Third, we conduct over-time analyses examining whether e-loadings exhibit systematic patterns as a function of firm age, where age proxies for the amount of information available about the firm. We predict that investor perceptions of earnings quality are more uncertain (leading to higher e-loadings) and less stable (leading to lower autocorrelation in e-loadings) for young firms where less information is available. Fourth, we examine whether e-loadings are higher in three settings associated with poor earnings quality: restatements, shareholder lawsuits, and bankruptcy.

In terms of the first analysis, we find that earnings quality determinants are significant in explaining variation in e-loadings, and that e-loadings exhibit predictably positive correlations with most other proxies for earnings quality. Our second analysis shows that firms with higher e-loadings have lower earnings response coefficients and more disperse and less accurate analysts' forecasts, consistent with market participants perceiving higher e-loading firms as having noisier earnings signals than lower e-loading firms. Results of our third analysis—of differences in the level and stability of e-loadings as a function of firm age—are also consistent with predictions: as the firm matures, we find both an over-time decline in the magnitude of its e-loading as well as an over-time increase in the autocorrelation of its e-loading. Our fourth analysis reveals that e-loadings are larger in settings characterized by poor earnings quality: e-loadings increase prior to, and are highest during, years containing restatement announcements, lawsuit filings, or bankruptcies. The level and change in e-loadings for firms affected by these events is also significantly larger than the levels and changes observed for samples of non-event firms, matched in calendar time. Results for all analyses are significant at the 0.05 level or better. We interpret the combined evidence as demonstrating the reliability of e-loadings as a returns-based representation of earnings quality, controlling for other factors known to affect returns.

To examine the robustness and generalizability of our findings, we extend our tests in several ways. First, we verify that our results are not driven by the subset of firms with the necessary data to calculate the accounting measure (*AQ*) that underlies *AQfactor*. Second, we examine whether e-loadings based on other measures of earnings quality perform as well as e-loadings based on accruals quality. Of the seven measures that we consider (persistence, predictability, smoothness, value relevance, timeliness, conservatism, and a measure of abnormal accruals), e-loadings based on mimicking factors for both persistence and smoothness appear to perform about as well as e-loadings based on accruals quality. However, none of the seven measures systematically dominates e-loadings based on accruals quality. Third, we probe the reliability of e-loadings calculated over *quarterly* estimation intervals, where we require at least 45 daily returns in a quarter to calculate the firm-quarter e-loading. Quarterly e-loadings exhibit the same patterns found for yearly e-loadings, except that the confidence intervals are wider; however, results are significant at the 0.10 level or better. Fourth, we examine the overlap between e-loadings and two trading-based measures of information asymmetry—bid-ask spreads and probability of informed trading (PIN) scores. The correlations range between 0.12 and 0.16 and, while reliably different from zero, are relatively weak in economic terms. To determine the influence of spreads and PIN scores on e-loadings, we orthogonalize e-loadings with respect to these measures and repeat our tests. Results are similar, suggesting that the effects we document for e-loadings are not subsumed or driven by trading-based measures of information asymmetry, such as bid-ask spreads and PIN scores.

We believe that the strength and consistency of our results, including results of extensions, alleviate concerns that arise because *AQfactor* itself is constructed using a sample of firms with the requisite time-series data to estimate the *AQ* measure. If this sample is not sufficiently broad, then the resulting returns representation of *AQ* demonstrated by *AQfactor* may not reflect the market's pricing of *AQ*, impairing our ability to identify a meaningful returns-based representation of earnings quality. In particular, if *AQfactor* is sufficiently noisy, then we would not expect it to load positively in asset-pricing regressions. Our findings that *AQfactor* exhibits significant (at the 0.001 level) average loadings for our sample and that *AQfactor* itself is associated with an economically substantial and statistically reliable (at the 0.001 level) average return indicate that whatever bias is introduced by using an accounting-based sample as the foundation for *AQfactor* does not invalidate the approach. Taken together, we believe the results support the inference that e-loadings capture the same underlying construct as reflected in other measures of earnings quality. That is, e-loadings capture what they purport to capture *and* offer several advantages, relative to accounting-based measures, in terms of flexibility and adaptability.[4]

While we believe that our results demonstrate that an earnings quality measure based on e-loadings offers significant advantages compared to existing earnings quality measures, our results do not speak to whether e-loadings are the *best* measure. The choice of the best earnings quality measure will be a function of, among other things, the nature of the research question addressed, the assumptions necessary to support the chosen research design, and available data. For example, the e-loading explored in this paper is a measure of investors' perception of *total* earnings quality, not discretionary earnings quality. In a research setting that demands a measure of the portion of earnings quality that is solely attributable to managements' discretionary actions and behaviors, a different measure of earnings quality would likely be preferred.

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[4] Our evidence supports Chen et al.'s (2004) analysis of over-time changes in the information risk characteristics of firms that change dividend policies. They find that information risk (proxied by the coefficient estimate on monthly *AQfactor*) is higher in the three years following a dividend decrease relative to the three years preceding the decrease; they find mixed results for dividend-increasing firms.

Two other caveats about the use of e-loadings are in order. First, unlike measures that rely exclusively on accounting data, e-loadings capture the market's *perception* of earnings quality. If market beliefs are not rational, then those perceptions may differ from the reality of the firm's earnings quality. (Because our tests focus on broad samples of firms over a 34-year period, it is unlikely that our results are attributable to systematic mispricing.) Second, our analyses of the reliability of e-loadings focus on the cross-sectional distribution of empirical measures of an unobservable construct, not on the correctness of an estimated magnitude.[5] Our results have no implications for the magnitudes of the e-loadings, and theory does not predict values for these magnitudes. In this respect, e-loadings are similar to *s*- and *h*-loadings (for *SMB* and *HML*); only the CAPM beta has a value that derives from theory.

The rest of the paper is organized as follows. The next section elaborates on the characteristics of returns-based measures of earnings quality and their advantages relative to existing earnings quality measures. Section III lays out the construction of the *AQ* factor-mimicking portfolio and the estimation of the e-loadings. Section IV describes the main samples used in our tests; Section V describes the empirical analyses; and Section VI presents extensions. Section VII summarizes the findings and concludes.

II. RETURNS-BASED REPRESENTATIONS VERSUS ACCOUNTING-BASED REPRESENTATIONS OF EARNINGS QUALITY

In this section we describe how a returns-based representation of earnings quality offers two distinct but related advantages over earnings quality measures derived from accounting data. (While we focus on accounting-based measures, much of the discussion also applies to measures derived from a combination of accounting and market data, for example, the explanatory power of earnings for returns.) Those advantages derive from differences in data requirements and periodicity of estimation. We also describe the types of research settings where a returns-based measure like the one we propose and validate might be used, and where it would not be useful.

Data Requirements

Measures of earnings quality based on accounting data are typically estimated using either a firm-specific time-series of annual data or industry cross-sections.[6] Either approach places significant restrictions on sample sizes and, in the case of firm-specific time-series estimation, biases the sample toward surviving firms that tend to be larger and more profitable. To illustrate this point, we refer to Dechow and Dichev (2002, Table 1), showing the derivation of their sample used to measure accruals quality as the mapping of current accruals into this year's, last year's, and next year's cash from operations (*CFO*). A key determinant of the sample size and composition is how many firms and of what type have data on current accruals and *CFO*. Dechow and Dichev (2002, Table 1) show that 55,850 firm-years between 1987–1999 have data on *CFO*, earnings, and changes in accounts receivable and inventory and total assets (after truncation of the most extreme 1 percent of observations). However, only 30,317 firm years have both lead and lag values of *CFO*—

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[5] This aspect of our analysis is similar to some aspects of Botosan and Plumlee's (2005) calibration of empirical proxies for the cost of equity capital. They focus primarily on the adequacy of cross-sectional distributions by examining associations between measures of firm-specific risk (such as beta) and measures of the cost of capital.

[6] An alternative to the use of annual data is to use quarterly accounting data; this approach introduces possible seasonal effects as well as sampling restrictions arising from availability of quarterly data.

reducing the sample size by nearly 46 percent—and 15,234 firm years (1,725 firms) have eight or more annual observations necessary to estimate firm-specific regressions—a reduction of nearly 73 percent relative to the original sample.

The alternate approach of estimating measures in industry cross-sections also places severe restrictions on sample size and composition. For example, Xie (2001) follows Subramanyam's (1996) cross-sectional modified Jones model estimation and reports a sample size of 56,692 firm-years over 1971–1992 for a relatively broad industry definition: at least six firms in a two-digit SIC code (he also excludes Nasdaq firms prior to 1982). Over this period, there are 106,645 firm-years with non-missing returns data on CRSP (also excluding Nasdaq firms prior to 1982), indicating a sample loss of about 47 percent. The decrease in sample size is more extreme if the industry requirements are tightened either by requiring more firms in the industry or by using more precise industry definitions. For example, Dechow and Dichev (2002, Table 1) report that, starting with 55,850 firm-year observations, 27,204 firm-year observations (136 three-digit industries) remain—a reduction of about 51 percent—after imposing the requirements of three years of *CFO* data and at least 50 observations per industry.

To summarize, we are not aware of accounting-based measures of earnings quality that do not impose data constraints that result in substantial decreases in sample size and (in the case of firm-specific time-series estimation) increases in survivorship bias. The decreases in sample size that arise in firm-specific time-series estimations are well known. Industry-estimation can be done using short series (e.g., as short as two or three years), but even this restriction reduces sample sizes and the requirement for a minimum number of firms per industry further restricts the sample. In contrast, a returns-based representation requires only a sufficient *returns* series. We show that, despite the additional noise introduced by daily returns, reliable estimates of e-loadings can be obtained for periods as short as one quarter. Thus, e-loadings can be calculated for larger and more representative (of the population) samples than can accounting-based measures of earnings quality. This is a distinct advantage in research settings where generalizability is important or where posited effects are expected to be most likely present in smaller, younger firms that lack the data necessary to calculate accounting-based measures.

Periodicity of Measurement

In extensions of our main findings (reported in Section VI), we show that e-loadings can be reliably calculated over intervals as short as 45 trading days, or roughly one quarter. Further, these 45 days need not be aligned with reporting periods. These two features mean that e-loadings can be used to examine shifts in earnings quality over short intervals and around events that occur at any time. In contrast, accounting-based measures of earnings quality are, by construction, linked to annual or quarterly reporting periods, cannot be applied to short intervals, and cannot be specific to a given financial statement date because they will be based on both current and prior data.

Research Settings Where Returns-Based Representations of Earnings Quality Offer Advantages

The two advantages of e-loadings (or other returns-based representations of earnings quality) discussed previously point to the places where these earnings quality measures can be particularly useful. Because e-loadings can be calculated for samples comprised of younger, smaller firms that have (at least) one or more quarters of daily returns data but lack a time-series of accounting data, research questions involving, for example, earnings quality for firms that have gone public within the last two years are more readily addressed using a returns-based representation.

While the sampling advantage means that using e-loadings increases sampling power and generalizability, the periodicity advantage means that e-loadings can be used when the research question of interest pertains to an event, occurring at any time, that has the potential to shift investor perceptions of earnings quality. Our reliability assessments in Section V, for example, focus on shifts in investor perceptions of earnings quality surrounding bankruptcies, restatements, and class action lawsuits. Other events where analysis of e-loadings could provide useful insights include adoptions (including early adoptions) of new accounting standards, voluntary accounting changes, and management or auditor changes that are prompted by concerns about financial reporting quality. In some settings, a shift in earnings quality might appear in conjunction with a shift in fundamental risks. Examples include changes in capital structure, mergers/acquisitions, and divisive restructurings (e.g., spin-offs that change the number and nature of segments and therefore the innate factors that influence earnings quality). In these settings, an analysis of shifts in information risk would have to control for the effects of shifts in fundamental risks.

Finally, the approach used to develop e-loadings is flexible with regard to other earnings attributes. That is, the factor-mimicking approach we use could be applied to attributes called for by a specific research question, such as smoothness or persistence. Therefore, if the research question requires, for example, a short-interval assessment of possible shifts in investor perceptions of earnings persistence, it would be possible to apply the approach we use to develop persistence-based e-loadings.

Research Settings Where Returns-Based Representations of Earnings Attributes Are Inapplicable

There are two kinds of settings where the approach we describe would be either inapplicable or perhaps of low power. With regard to the issue of power, recall that the e-loading measure captures investor perceptions of *total* earnings quality; however, in some settings it is *discretionary* earnings quality that is of interest. Previous research (Francis et al. 2005) shows that the portion of accruals quality that is due to management's reporting choices is less priced than is total or innate accruals quality. This distinction has implications for the power of tests requiring a measure of the discretion in management's short-term decisions.

With regard to the question of applicability, a returns-based representation of an earnings attribute cannot capture the sign of the underlying factor (e.g., positive or negative abnormal accruals). A returns-based representation would not, therefore, be applicable in settings that require tests of directional predictions concerning opportunistic earnings management (e.g., tests of whether earnings are managed up or down in response to some posited incentive).

III. *AQ* AND *AQ* FACTOR-MIMICKING PORTFOLIOS

Our main tests use Dechow and Dichev's (2002) (hereafter, Dechow and Dichev) accruals quality metric to capture earnings quality. This choice is based on Francis et al.'s (2004) finding that accruals quality has a stronger association with the cost of equity than other earnings attributes. As discussed in Francis et al. (2005), theories developed by Easley and O'Hara (2004) and Leuz and Verrecchia (2005) predict that information risk is priced by investors. Positing that investors price securities based on information about cash flows, Francis et al. (2005) use accruals quality (which captures the imprecision of the mapping between current accruals and cash flows) as a measure of information risk.

We operationalize accruals quality using McNichols' (2002) modification of Dechow and Dichev's model:

$$TCA_{j,T} = \phi_{0,j} + \phi_{1,j}CFO_{j,T-1} + \phi_{2,j}CFO_{j,T} + \phi_{3,j}CFO_{j,T+1} + \phi_{4,j}\Delta \mathrm{Rev}_{j,T} + \phi_{5,j}PPE_{j,T} + \upsilon_{j,T} \quad (1)$$

where $TCA_{j,T} = \Delta CA_{j,T} - \Delta CL_{j,T} - \Delta Cash_{j,T} + \Delta STDEBT_{j,T}$ = total current accruals in year T; $CFO_{j,T} = NIBE_{j,T} - TA_{j,T}$ = firm j's cash flow from operations in year T;[7] $NIBE_{j,T}$ = firm j's net income before extraordinary items (Compustat #18) in year T; $TA_{j,T} = (\Delta CA_{j,T} - \Delta CL_{j,T} - \Delta Cash_{j,T} + \Delta STDEBT_{j,T} - DEPN_{j,T})$ = firm j's total accruals in year T; $\Delta CA_{j,T}$ = firm j's change in current assets (Compustat #4) between year $T-1$ and year T; $\Delta CL_{j,T}$ = firm j's change in current liabilities (Compustat #5) between year $T-1$ and year T; $\Delta Cash_{j,T}$ = firm j's change in cash (Compustat #1) between year $T-1$ and year T; $\Delta STDEBT_{j,T}$ = firm j's change in debt in current liabilities (Compustat #34) between year $T-1$ and year T; $DEPN_{j,T}$ = firm j's depreciation and amortization expense (Compustat #14) in year T; $\Delta Rev_{j,T}$ = firm j's change in revenues (Compustat #12) between year $T-1$ and year T; and $PPE_{j,T}$ = firm j's gross value of property, plant, and equipment (Compustat #7) in year T. All variables are scaled by average assets.

We estimate Equation (1) in annual industry cross-sections for each of the 48 Fama and French (1997) industries with at least 20 firms in that industry-year. These estimations produce firm-year residuals, $\hat{\upsilon}_{j,T}$. Our accounting-based earnings quality metric for firm j in year T is the standard deviation of firm j's residuals over the past five years, $AQ_{j,T} = \sigma(\hat{\upsilon}_{j,T})$, $T = T-5,...,T-1$.[8] Large (small) values of $AQ_{j,T}$ correspond to poor (good) accruals quality.

We assign firms to *AQ* deciles using a dynamic portfolio technique that allows for differences in firms' fiscal year-ends as well as over-time changes in accruals quality. Specifically, we form deciles on the first day of each month m based on the firm's most recent value of *AQ* known prior to m; firms with the smallest (largest) *AQ* values are placed in the first (tenth) decile. This means that firm j's *AQ* signal for fiscal year T, where fiscal year T ends in month n, will influence firm j's ranking for months $n+4$ through $n+15$. We then calculate the average daily return for each decile for the period January 2, 1970 (the first trading day of 1970) to December 31, 2003, yielding a time-series of 8,586 daily returns for each decile (D1,...,D10). The *AQ* factor-mimicking portfolio, *AQfactor*, equals the difference between the daily returns of the poorest *AQ* deciles (deciles 7–10) and the best *AQ* deciles (deciles 1–4). This procedure (similar to that used by Carhart [1997] to construct a factor-mimicking portfolio for price momentum) yields a series of 8,586 daily *AQfactor* returns ($AQfactor_t$).[9]

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[7] We calculate total accruals using information from the balance sheet and income statement rather than from the statement of cash flows (as advocated by Hribar and Collins [2002]) because statement of cash flow data are not available prior to 1988 (the effective year of SFAS No. 95) and our *AQ* metric requires seven yearly observations.

[8] Calculating the *AQ* measure in year T using the residuals in years $T-5$ to $T-1$ accounts for the fact that Equation (1) contains a lead term, $CFO_{j,T+1}$. In total, calculation of *AQ* requires seven years of data because we require five residuals and the model contains two lag terms.

[9] Carhart's (1997) procedure differs from Fama and French's (1993) procedures used to create size and book-to-market mimicking factors in two respects: he uses equal-weighted rather than value-weighting and does not orthogonalize with respect to firm size. As sensitivity tests, we recreate an *AQfactor* that is value-weighted within each cell and orthogonalized with respect to size. (Both value-weighting and orthogonalization serve to remove any size-shared variation from the estimation of the e-loadings.) Results based on the value-weighted, size-orthogonalized *AQfactor* (not reported) are similar in all respects to those documented.

For our sample, the average value of $AQfactor_t$ is 0.0779 percent, or about 22 percent on an annualized basis; the standard error is 0.0085 percent. To put these figures in perspective, over the same time period, the average daily excess market return is 0.0216 percent (standard error of 0.0102 percent), the average value of the daily *SMB* is 0.0021 percent (standard error of 0.0056 percent), and the average value of the daily *HML* is 0.0224 percent (standard error of 0.0052 percent). *AQfactor* is reliably different from zero (t-statistic = 9.17, or the mean value of 0.0779 percent divided by its standard error of 0.0085 percent), and has a higher t-statistic than the other risk factors (the t-statistics for market risk, *SMB*, and *HML* are 2.12, 0.37, and 4.28, respectively).

Because *AQfactor* is time-specific, not firm-specific, we can correlate *AQfactor* with the returns of *any* firm to determine that firm's exposure to poor earnings quality, much like we correlate a firm's returns with the market risk premium to obtain a measure of its exposure to market risk. The specific correlation measure we use is the coefficient estimate on *AQfactor* obtained from one-factor (superscript $1f$) and three-factor (superscript $3f$) asset-pricing regressions that include *AQfactor* as an independent variable:

$$R_{j,t} - R_{F,t} = \alpha^{1f}_{j,T} + \beta^{1f}_{j,T}(R_{M,t} - R_{F,t}) + e^{1f}_{j,T}AQfactor_t + \varepsilon^{1f}_{j,t} \tag{2}$$

$$R_{j,t} - R_{F,t} = \alpha^{3f}_{j,T} + \beta^{3f}_{j,T}(R_{M,t} - R_{F,t}) + s^{3f}_{j,T}SMB_t + h^{3f}_{j,T}HML_t + e^{3f}_{j,T}AQfactor_t + \varepsilon^{3f}_{j,t} \tag{3}$$

where t = index for the number of trading days in year T; $R_{j,t}$ = firm j's return on day t; $R_{F,t}$ = the risk-free rate on day t; $R_{M,t}$ = the market return on day t; SMB_t = small-minus-big factor on day t; HML_t = high-minus-low book-to-market factor on day t. $R_{M,t}$, SMB_t and HML_t are obtained from the WRDS database (where raw returns are from CRSP and *SMB* and *HML* factors are from Ken French). For the one-factor (three-factor) specification, $e^{1f}_{j,T}$ ($e^{3f}_{j,T}$) is the estimate of firm j's sensitivity to poor earnings quality in year T. The other slope coefficients, $\beta^{1f}_{j,T}$ (or $\beta^{3f}_{j,T}$), $s^{3f}_{j,T}$, and $h^{3f}_{j,T}$, capture the firm's exposure to returns representations of market risk, size, and book-to-market, respectively, in year T.

In summary, we create an accounting-based measure of accruals quality $AQ_{j,T}$ using a restricted sample of firms with seven years of accounting data, convert *AQ* to a time-specific returns representation ($AQfactor_t$), and use this time-specific returns representation in firm- and year-specific regressions to obtain a firm-year returns-based representation of earnings quality ($e^{\#}_{j,T}$, $\# = 1f$ or $3f$). Relative to the original accounting-based measure, e-loadings can be calculated for much larger samples because they require only enough daily returns in year T to estimate Equations (2) or (3). An additional benefit of the returns approach is that because $AQfactor_t$ varies through time, e-loadings are not constrained to be slow to change, as is *AQ*, which requires seven years of accounting data and, therefore, has a mechanical dependence year over year.

IV. SAMPLE AND DESCRIPTIVE DATA

We begin by identifying all firms with the necessary data to estimate *AQ* in each year T = 1970–2003 (the AQ Sample).[10] Using the *AQ* measures, we calculate *AQfactor* using the procedures described in Section III. Next, we identify all firms with at least 100 daily returns in year *T* (the Returns Sample). The requirement of 100 daily returns in year *T* to estimate Equations (2) and (3) is *ad hoc*, and we assess its sensitivity in Section VI. Table 1, Panel A shows the number and percentage of sample firms, by year, for the two samples. As expected given the sample selection criteria, the Returns Sample is larger than the AQ Sample, both in number of firms (mean of 6,408 firms per year versus 2,147 firms per year) and as a percentage of traded firms (mean of 92.4 percent per year versus 30.9 percent). To further contrast the representativeness of the two samples, Panel B reports information about the distributions of size (market capitalization, total assets, and total sales) and performance (as measured by return on assets, ROA, and return on equity, ROE). Comparisons show that the AQ Sample contains larger firms (mean market capitalization is \$1,295 million versus \$942 million; mean assets are \$1,404 million versus \$1,005 million; mean sales are \$1,298 million versus \$908 million) and more successful firms (mean ROA is 3.7 percent versus 2.6 percent; mean ROE is 7.3 percent versus 5.5 percent). These data demonstrate that the Returns Sample dominates the AQ Sample on both sample size and survivorship bias. Given that the Returns Sample is minimally restricted (i.e., it requires only 100 daily returns in year *T*), it is more representative of the population than is the AQ Sample, which requires a firm to have seven years of accounting data for inclusion in year *T*.

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[10] Because of the data requirements for *AQ*, as well as the additional three-month lag we require for accounting data to reach the market, we use accounting data from as early as 1963.

We next estimate firm-year e-loadings for the Returns Sample, using Equations (2) and (3); we discuss but do not tabulate these results. For e-loadings based on the CAPM, the mean value of e^{1f} is 0.3021 with a standard deviation of 0.5335, and the inter-quartile range is −0.0089 (25th percentile) to 0.5249 (75th percentile). For e-loadings based on the three-factor model, the mean value of e^{3f} is 0.1154 with a standard deviation of 0.5860, and the inter-quartile range is −0.1694 (25th percentile) to 0.3126 (75th percentile). The other loadings in these models (i.e., β for the CAPM and β, s, and h for the three-factor model) exhibit similar variation. The mean and median explained variation of the augmented one-factor and three-factor models is between 3 and 9 percent, less than the 16–20 percent reported by Francis et al. (2005) for estimations requiring a minimum of 18 monthly observations per firm. The lower explained variability is expected because we estimate our models using daily returns, which are noisier than monthly returns.[11]

V. ANALYSIS OF E-LOADINGS

Our four analyses of the reliability of e-loadings are complementary, in that they use different research designs and distinct indicators of earnings quality. We begin by using a cross-sectional design to analyze the associations between e-loadings and both determinants of accruals quality and measures of earnings quality that have been used in other research settings. Our second set of tests also uses a cross-sectional design, but focuses on earnings response coefficients and properties of analyst forecasts. The third test shifts to a time-series design, and considers the over-time relation between e-loadings and firm age as a summary indicator of available firm-specific information. Finally, we analyze the relation between e-loadings and events associated with poor earnings quality.

Associations between e-Loadings and Determinants of and Proxies for Earnings Quality

Our first analysis considers whether accounting-based determinants of earnings quality explain cross-sectional variation in e-loadings. We focus on the five innate determinants

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[11] Note that these R^2s are based on firm-specific estimations; calendar-time portfolio regressions generally show much higher explanatory power (often above 40 percent) because firm-specific variation is largely eliminated in the portfolio design.

TABLE 1
Descriptive Information on AQ Sample and Returns Sample[a]

Panel A: Firm Representation in AQ Sample and Returns Sample

	AQ Sample		Returns Sample			AQ Sample		Returns Sample	
Year	# Firms	% Traded Firms	# Firms	% Traded Firms	Year	# Firms	% Traded Firms	# Firms	% Traded Firms
1970	708	28.4	2,405	96.3	1988	2,194	28.4	7,228	93.4
1971	790	30.3	2,496	95.6	1989	2,155	28.9	7,006	93.9
1972	873	15.2	2,648	46.0	1990	2,249	30.8	6,951	95.3
1973	949	16.0	5,634	94.7	1991	2,375	32.4	6,830	93.2
1974	1,047	19.5	5,191	96.5	1992	2,436	31.9	7,067	92.7
1975	1,197	23.1	5,005	96.4	1993	2,583	31.8	7,455	91.6
1976	1,336	25.4	5,048	96.1	1994	2,779	31.9	8,225	94.5
1977	1,444	27.6	5,057	96.6	1995	2,792	30.7	8,431	92.7
1978	1,480	28.5	4,941	95.1	1996	2,785	28.7	8,967	92.5
1979	1,865	36.3	4,895	95.2	1997	2,773	27.9	9,298	93.6
1980	2,381	44.7	4,939	92.7	1998	2,758	28.1	9,251	94.3
1981	2,673	46.8	5,370	94.0	1999	2,703	28.5	8,703	91.8
1982	2,599	43.6	5,448	91.4	2000	2,751	30.0	8,571	93.3
1983	2,545	38.6	5,919	89.8	2001	2,803	33.1	7,889	93.2
1984	2,436	35.5	6,472	94.3	2002	2,880	37.1	7,331	94.3
1985	2,336	33.4	6,454	92.4	2003	2,920	39.8	6,877	93.8
1986	2,229	30.2	6,661	90.3	Mean	2,147	30.9	6,408	92.4
1987	2,168	28.1	7,219	93.6	Median	2,378	30.2	6,746	93.7

Panel B: Descriptive Statistics for Selected Financial Variables

	AQ Sample						
	Mean	Std. Dev.	10%	25%	Median	75%	90%
Market value	1294.97	4938.84	7.97	24.95	114.08	617.69	2539.33
Total assets	1403.51	4123.69	15.53	43.90	170.94	800.05	3137.31
Sales	1298.30	3451.51	16.99	53.93	210.98	872.37	3109.41
ROA	0.037	0.086	−0.043	0.015	0.046	0.078	0.114
ROE	0.073	0.183	−0.106	0.035	0.105	0.157	0.213

	Returns Sample						
	Mean	Std. Dev.	10%	25%	Median	75%	90%
Market value	942.49	4133.90	6.01	17.44	72.41	375.47	1588.85
Total assets	1004.68	3488.89	10.19	27.51	98.75	443.31	1985.70
Sales	907.51	2874.54	9.60	30.40	114.79	483.29	1896.36
ROA	0.026	0.106	−0.078	0.008	0.044	0.078	0.118
ROE	0.055	0.208	−0.179	0.019	0.100	0.156	0.218

Sample descriptions: The AQ Sample consists of all firms with at least seven years of time-series Compustat data on working capital accruals and lead, lag, and current cash flows to estimate Dechow-Dichev regressions. The Returns Sample includes all firms with at least 100 CRSP daily returns in year *T*.

[a] Panel A reports the distribution of the AQ Sample and, separately, the Returns Sample by year. Panel B reports summary financial data on the size and profitability of the firms in each sample. Market value, total assets, and sales are in millions of dollars. Variables are measured over the entire sample period 1970–2003.

proposed by Dechow and Dichev: firm size (*log(assets)*, measured as the log of total assets; results are not sensitive to other measures of size, such as revenues), standard deviation of cash flow from operations ($\sigma(CFO)$), standard deviation of sales revenues ($\sigma(Sales)$), length of operating cycle (*OperCycle*, measured as the log of the sum of days accounts receivable and days inventory), and negative earnings incidence (*NegEarn* = 1 if year *T* earnings are negative, 0 otherwise). We measure *log(assets)*, *log(OperCycle)*, and *NegEarn* as of the end of year *T* and we calculate $\sigma(CFO)$ and $\sigma(Sales)$ using data from year $T-6$ to *T*. Because of the latter calculation, the sample used for these tests is the AQ Sample, not the Returns Sample.

Table 2, Panel A shows the results of regressing the year *T* e-loadings on the innate determinants. We report the average values of the coefficient estimates obtained from 34 yearly regressions; t-statistics are based on the standard errors of the 34 annual coefficients (Fama and MacBeth 1973). The results indicate that e-loadings are highly correlated with each of the innate determinants, in the directions predicted by Dechow and Dichev: e-loadings are negatively correlated with firm size, and positively correlated with the variability of cash flows and sales, the length of the operating cycle, and the incidence of losses (t-statistics range in absolute value from 8.70 to 16.10).

Our examination of the relation between e-loadings and other measures of earnings quality focuses on the seven earnings attributes considered by Francis et al. (2004): accruals quality itself (*AQ*), earnings persistence (*Persistence*, measured as the negative of the AR1 parameter from firm-specific regressions of current earnings per share on lagged earnings per share), earnings predictability (*Predictability*, measured as the standard deviation of the error term from firm-specific AR1 models of earnings), smoothness (*Smoothness*, measured as the ratio of the standard deviation of earnings to the standard deviation of cash flows), value relevance (*Value Relevance*, measured as the negative of the explained variability of a regression of annual returns on the level and change in earnings per share), timeliness (*Timeliness*, measured as the negative of the explained variability from a Basu [1997] reverse regression of earnings on returns controlling for the sign of those returns), and conservatism (*Conservatism*, measured as the negative of the coefficient on negative returns from the aforementioned reverse regression). We follow Francis et al.'s (2004) procedures and estimate each attribute over rolling ten-year windows. Associations are calculated between these measures of earnings quality and the e-loadings averaged across the same rolling ten-year windows. Each measure is ordered consistently such that higher (lower) values indicate poorer (better) earnings quality.

Table 2, Panel B shows the pairwise correlations between the e-loadings and the earnings quality proxies. To the extent that the e-loadings are correlated with the constructs captured by each of these attributes, we expect to observe positive correlations between e-loadings and each variable. For example, the correlation between e-loadings and *AQ* provides evidence on whether e-loadings capture the variable that underlies the construction of the *AQfactor*. We expect and find that this correlation is positive, with magnitudes ranging from 0.4397 to 0.5008, all significant at the 0.0001 level.[12] Francis et al.'s (2004) results suggest that associations between the e-loadings and the earnings quality attributes should be strong for the three attributes based only on accounting numbers, and weak for the three attributes that are based also on returns. Our results are for the most part consistent with their findings. Both *Persistence* and *Smoothness* exhibit significant (at the 0.0001 level)

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[12] This result is similar to Fama and French's (1993) evidence that loadings on the *SMB* and *HML* mimicking factors are positively correlated with their underlying variables, market capitalization and book-to-market, respectively.

TABLE 2
Associations between e-Loadings and Innate Determinants of Earnings Quality and Other Proxies for Earnings Quality

Panel A: Regressions of e-Loadings on Innate Firm Characteristics[a]

	Base Model: CAPM		Base Model: 3-Factor Model	
	Coefficient	**t-statistic**	**Coefficient**	**t-statistic**
log(Assets)	−0.0584	−15.86	−0.0339	−13.99
$\sigma(CFO)$	0.7026	13.54	0.6550	11.86
$\sigma(Sales)$	0.2304	16.10	0.1707	11.51
log(OperCycle)	0.0539	13.62	0.0383	12.13
NegEarn	0.1248	8.70	0.1490	13.55
Adj. R^2	0.2108		0.1154	

Panel B: Correlations between e-Loadings and Other Proxies for Earnings Quality[b]

	Base Model: CAPM		Base Model: 3-Factor Model	
	Pearson	**Spearman**	**Pearson**	**Spearman**
Accruals Quality	0.4397	0.5008	0.4459	0.4918
p-value	<.0001	<.0001	<.0001	<.0001
Persistence	0.1653	0.1872	0.2148	0.2575
p-value	<.0001	<.0001	<.0001	<.0001
Predictability	0.0207	0.0120	0.0102	0.0760
p-value	<.0001	0.0047	0.0162	<.0001
Smoothness	0.2226	0.2023	0.2457	0.2337
p-value	<.0001	<.0001	<.0001	<.0001
Value Relevance	0.0088	0.0245	0.0416	0.0278
p-value	0.0379	<.0001	<.0001	<.0001
Timeliness	0.0264	0.0383	0.0138	0.0007
p-value	<.0001	<.0001	0.0011	0.8701
Conservatism	0.0081	0.0137	0.0034	0.0197
p-value	0.0553	0.0013	0.4234	<.0001

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[a] Panel A reports the results of annual regressions of e-loadings (based on the CAPM and, separately, 3-factor model) on variables capturing innate features of earnings quality: *log(Assets)* = firm *j*'s log of total assets in year *T* (a measure of size), $\sigma(CFO)$ = standard deviation of cash flows over year *T* through $T-6$; = $\sigma(Sales)$ = standard deviation of sales calculated over years $T-6$ to *T*; *log(OperCycle)* = log of the sum of days receivables and days inventory); and *NegEarn* = 1 if year *T* earnings are negative, 0 otherwise. We report the mean value of the 34 annual coefficients; t-statistics are based on the standard errors of the 34 annual coefficients.

[b] Panel B reports the correlation between e-loadings and several measures of earnings quality: Accruals Quality (*AQ*), Persistence (AR1 parameter from firm-specific regressions of current EPS on lagged EPS), Predictability (standard deviation of the residuals from firm-specific AR1 models of EPS), Smoothness (ratio of the standard deviation of earnings to the standard deviation of cash flows), Value Relevance (explained variability of a regression of annual returns on the level and change in earnings per share), Timeliness (explained variability from a regression of earnings on returns controlling for the sign of returns), and Conservatism (coefficient on negative returns from the aforementioned reverse regression). These proxies are estimated using rolling ten-year windows, and are ordered such that larger values denote poorer quality (i.e., we use the negatives of the calculated measures for Persistence, Value Relevance, Timeliness, and Conservatism). To compare these ten-year based measures of earnings quality with our firm- and year-specific e-loadings, we average the e-loadings across the same rolling ten-year windows.

positive correlations with e-loadings ranging in magnitude from 0.1653 to 0.2575; for *Predictability*, the correlations are lower (0.0102 to 0.0760; significant at the 0.02 level or better). Results for measures of earnings quality that also rely on returns data (*Value Relevance*, *Timeliness*, and *Conservatism*) are generally low, with correlations ranging from 0.0007 (insignificant) to 0.0416 ($p < 0.0001$).

Market Participant Behaviors as a Function of e-Loadings

In this section, we examine whether e-loadings are associated with predictable variation in investor and analyst behaviors. Our first analysis builds on prior research (e.g., Imhoff and Lobo 1992) that posits information uncertainty as a determinant of investors' response to earnings as captured by the coefficient relating returns to earnings (the earnings response coefficient, or ERC). Our test of whether firms with higher e-loadings have smaller ERCs is based on the following regression:

$$CAR(-1,0)_{j,t} = \lambda_0 + \lambda_1 UE_{j,t} + \lambda_2 UE_{j,t} \times e^{\#}_{j,T} + \zeta_{j,t} \tag{4}$$

where $CAR(-1,0)_{j,t}$ = firm j's two-day cumulative abnormal return over the quarterly earnings announcement, where abnormal return is defined as the raw return less the value-weighted market return; and $UE_{j,t}$ = unexpected earnings conveyed by firm j's quarterly earnings announcement made on day t, equal to firm j's reported earnings for quarter q less the consensus analyst forecast, scaled by firm j's share price 20 days before the earnings announcement date.

Results of estimating Equation (4) are shown in Table 3, Panel A. For these tests, unexpected earnings are calculated using analyst forecast data from Zacks, 1983–2002; on average, there are 2,885 firms per quarter. We report mean values of the coefficients obtained from the 80 quarterly estimations of Equation (4) over the period 1983–2002; the t-statistics are based on the time-series of the standard errors of the 80 coefficient estimates.[13] The results show that earnings news announced by higher e-loading firms is associated with a significantly weaker market response than is earnings news announced by lower e-loading firms: $\lambda_2 = -0.0415$ (t-statistic = -3.96) using e^{1f}, and $\lambda_2 = -0.0370$ (t-statistic = -4.32) using e^{3f}. We also estimate quarterly regressions that include *UE* interacted with other variables known to affect earnings response coefficients: whether the firm reported a loss in quarter q (*NegEarn* = 1 if reported earnings in quarter q are negative, 0 otherwise), firm size (*ln(Size)* = log of firm j's sales revenues), firm j's market-to-book ratio (*MB*), and the ratio of firm j's debt to equity (*Leverage*):

$$\begin{aligned} CAR(-1,0)_{j,t} = {} & \lambda_0 + \lambda_1 UE_{j,t} + \lambda_2 UE_{j,t} \times e^{\#}_{j,T} + \lambda_3 UE_{j,t} \times NegEarn_{j,q} + \lambda_4 UE_{j,t} \\ & \times \ln(Size)_{j,q} + \lambda_5 UE_{j,t} \times MB_{j,q} + \lambda_6 UE_{j,t} \times Leverage_{j,q} + \xi_{j,t}. \end{aligned} \tag{5}$$

Results of estimating Equation (5), reported in Panel B, continue to show that firms with larger e-loadings have smaller responses to earnings news (t-statistics for λ_2 are -2.16 using e^{1f} and -3.14 using e^{3f}).

Our second analysis of the relation between e-loadings and market participant behavior examines the dispersion and accuracy of analysts' earnings forecasts. Based on prior research, we predict that firms with higher e-loadings have more difficult-to-predict earnings,

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[13] An alternate approach is to estimate Equation (4) as a pooled regression, with t-statistics based on Newey and West (1987) adjusted standard errors to control for heteroscedasticity and autocorrelation. Inferences (not tabulated) are the same: in particular, the t-statistic for λ_2 is -7.17 (-7.61) for e^{1f} (e^{3f}).

TABLE 3
Association between e-Loadings and Investor Responses to Unexpected Earnings

Independent Variable	Base Model: CAPM		Base Model: 3-Factor Model	
	Coefficient	t-statistic	Coefficient	t-statistic
Panel A: Regression of Abnormal Returns on Unexpected Earnings News, Conditional on e-Loadings[a]				
UE	0.1575	10.20	0.1487	10.78
UE × *e-loading*	−0.0415	−3.96	−0.0370	−4.32
Panel B: Regression of Abnormal Returns on Unexpected Earnings, Conditional on e-Loadings and Control Variables[a]				
UE	0.5512	14.55	0.5457	14.84
UE × *e-loading*	−0.0219	−2.16	−0.0239	−3.14
UE × *NegEarn*	−0.3989	−13.72	−0.3976	−13.76
UE × *ln(Size)*	−0.0023	−0.75	−0.0017	−0.61
UE × *MB*	0.0147	3.77	0.0142	3.89
UE × *Lev*	−0.1155	−4.07	−0.1163	−4.13

[a] This table reports the mean coefficient estimates obtained from 80 quarterly estimations of two-day market reactions to earnings announcements on unexpected earnings, unexpected earnings interacted with e-loadings, and other variables known to affect the market reaction. *UE* = unexpected earnings conveyed in quarterly earnings announcement, measured using analysts' consensus forecast as the measure of expected earnings. Panel A shows the results of regressing cumulative abnormal returns in days (−1,0) on *UE* and *UE* interacted with e-loadings. Panel B shows similar regressions that include *UE* interacted with other variables known to affect the earnings response coefficient: negative reported earnings (*NegEarn* = 1 if reported earnings in quarter *q* are negative, 0 otherwise); firm size (*ln(Size)* = log of sales revenues); market-to-book ratio (*MB*); and the ratio of debt to equity (*Leverage*). t-statistics are based on the standard errors of the 80 quarterly coefficients.

resulting in both more dispersed and less accurate forecasts relative to lower e-loading firms. More specifically, prior research reports positive relations between random walk measures of earnings surprise—which we interpret as indicators of difficult-to-predict earnings—and both forecast dispersion and forecast accuracy.[14]

Following prior literature, we measure forecast dispersion as the standard deviation of analysts' EPS forecasts for quarter q, scaled by share price at the beginning of the quarter. Tests of dispersion are conducted on all firms with at least three earnings forecasts for quarter q issued in the three months preceding the announcement of quarter q earnings; the three or more forecasts produce a standard deviation of the forecasts made in quarter q for firm j, $Dispersion_{j,q}$. We measure forecast accuracy as the absolute forecast error (the difference between reported and forecasted EPS for quarter q, scaled by share price at the beginning of the quarter). Tests of forecast accuracy are conducted on all firms with at least one quarterly earnings forecast for quarter q in the three-month period preceding the quarter q earnings announcement; our measure of forecast accuracy, $|FE|_{j,q}$, is the average of the absolute forecast errors across all forecasts made about firm j for quarter q. Larger values of $Dispersion_{j,q}$ and $|FE|_{j,q}$ indicate more dispersed and less accurate forecasts, respectively. Similar to Table 3, these tests are conducted using analyst forecast data from Zacks, 1983–2002.

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[14] Lang and Lundholm (1996) and Bowen et al. (2002) measure earnings surprise as the absolute value of the difference between reported earnings for quarter q and reported earnings for quarter $q-4$, scaled by share price ($|UE_{j,q}^{SRW}|$). In addition to $|UE_{j,q}^{SRW}|$, Heflin et al. (2003) include two indicator variables to capture earnings surprise: the first equals 1 if the sign of the unexpected earnings news in the quarter q earnings announcement is negative ($NegUE_{j,q}^{SRW} = 1$ if $UE_{j,q}^{SRW} < 0$), the second equals 1 if reported earnings for quarter q are negative, and 0 otherwise ($Loss_{j,q}$).

Our tests of whether firms with greater forecast dispersion and larger absolute forecast errors have higher e-loadings are based on the following regressions:

$$Dispersion_{j,q} = \kappa_0 + \kappa_1 e^{\#}_{j,T} + \kappa_2 Age_{j,q} + \kappa_3 ln(Sales)_{j,q} + \zeta_{j,q}. \tag{6}$$

$$|FE|_{j,q} = \delta_0 + \delta_1 e^{\#}_{j,T} + \delta_2 Age_{j,q} + \delta_3 ln(Sales)_{j,q} + \psi_{j,q}. \tag{7}$$

Equations (6) and (7) control for forecast age and firm size, both of which have been shown to affect forecast dispersion and accuracy (Bowen et al. 2002). Forecast age ($Age_{j,q}$) is the average number of days that the forecasts issued in quarter q precede the earnings announcement date. Firm size is proxied by the log of sales in quarter q, $\ln(Sales)_{j,q}$. We estimate Equations (6) and (7) using both pooled and quarterly estimations; results are similar for the two approaches, so we tabulate only the quarterly results. Table 4 reports the mean values of the coefficients from the 80 quarterly regressions, with t-statistics based on the standard errors of those 80 estimates. The results for both dispersion (Panel A) and absolute forecast errors (Panel B) indicate that e-loadings are positively associated with both forecast properties (t-statistics range between 5.39 and 12.07). We also re-estimate Equations (6) and (7) including the three measures of earnings surprise considered in prior research (described in footnote 14). Even controlling for these other surprise measures, the results show that e-loadings remain significant (t-statistics range from 2.92 to 8.40) in explaining properties of analysts' forecasts.

Patterns in e-Loadings Associated with Firm Age

This section examines patterns in e-loadings associated with firm age, which we view as an inverse measure of investor perceptions of the amount and stability of firm-specific information. Our examination is predicated on the view that the relatively sparse firm-specific information that is available for younger firms induces greater uncertainty about that firm's business, including uncertainty about its financial reporting quality. We conduct two investigations involving firm age and e-loadings. The first involves newly listed firms; for such firms, we expect information uncertainty is highest at initial listing and declines as more information becomes available as the firm matures. This hypothesis follows from Lang's (1991) argument (in an ERC setting and in the context of IPO firms) that, as more earnings observations are revealed, investors' uncertainty about the persistence parameter underlying the earnings process decreases.[15] Similar to Pastor and Veronesi (2003) and Fama and French (2004), we measure firm age as the number of years since listing (*FirmAge*), where listing year (year 0) is the first year for which stock returns are available on CRSP. We consider age, rather than measures such as firm size, institutional holdings, and analyst coverage (which have also been used to reflect the relative uncertainty or quality of information) so that we can use an over-time design to examine whether e-loadings decline after initial stock market listing.

Figure 1 shows the trend in mean e-loadings as a function of *FirmAge* for firms in the Returns Sample that listed in any year, 1926–2003. To be included in Figure 1, the firm

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[15] Regulation S-X requires IPO firms to provide three years of income and cash flows in their Form S-1 registration statements. Firms in existence less than three years must provide information for the period they have been in existence. Thus, for most newly listed firms, financial data exist that provide a basis for investors to develop views about the firm's financial reporting quality.

TABLE 4
Relation between e-Loadings and Forecast Dispersion and Forecast Accuracy

	Base Model: CAPM				**Base Model: 3-Factor**			
Variable	**Coef.**	**t-stat**	**Coef.**	**t-stat**	**Coef.**	**t-stat**	**Coef.**	**t-stat**
Panel A: Regression of Forecast Dispersion on e-Loadings and Control Variables[a]								
e-loading	0.00159	7.99	0.00080	6.31	0.00160	12.07	0.00081	8.40
Age	−0.00001	−6.34	−0.00001	−6.55	−0.00001	−6.1	−0.00001	−6.47
ln(size)	0.00001	0.22	0.00006	3.03	−0.00011	−7.12	0.00001	0.57
AbsUE	NA	NA	0.04593	8.82	NA	NA	0.04624	8.79
NegUE	NA	NA	0.00073	13.37	NA	NA	0.00072	12.88
Loss	NA	NA	0.00360	17.76	NA	NA	0.00353	17.07
Adj. R^2	0.0337		0.2835		0.0374		0.2848	
Panel B: Regression of Forecast Accuracy on e-Loadings and Control Variables[a]								
e-loading	0.00400	5.62	0.00138	2.92	0.00352	5.39	0.00137	3.49
Age	0.00003	6.25	0.00004	7.84	0.00003	6.25	0.00004	7.86
ln(size)	0.00065	−7.57	0.00040	4.93	−0.00081	−9.64	0.00037	4.62
AbsUE	NA	NA	0.39921	10.68	NA	NA	0.39775	10.67
NegUE	NA	NA	0.00436	11.05	NA	NA	0.00436	11.13
Loss	NA	NA	0.01718	19.27	NA	NA	0.01730	18.60
Adj. R^2	0.0171		0.4540		0.0168		0.4534	

Sample description and variable definitions: The tests in this table require data on both e-loadings and analyst forecasts. For tests of forecast dispersion (Panel A), we require the firm to have at least three forecasts for quarter q. For tests of forecast accuracy (Panel B), we require the firm to have at least one forecast for quarter q. $Dispersion_{j,q}$ = the price-scaled standard deviation of analysts' earnings forecasts for firm j's quarter q made in the three months preceding the quarter q earnings announcement; $|FE|_{j,q}$ = the average value of the price-scaled absolute forecast error for all quarter q forecasts about firm j made in the three months preceding the quarter q earnings announcement; $Age_{j,q}$ = the average number of days between the forecast and the earnings announcement date; $ln(Sales)_{j,q}$ = log of firm j's sales for quarter q; $|UE^{SRW}_{j,q}|$ = the absolute of the seasonal random walk (SRW) measure of the unexpected earnings revealed in the quarter q earnings announcement, scaled by share price; $NegUE^{SRW}_{j,q} = 1$ if $UE^{SRW}_{j,q} < 0$, 0 otherwise; and $Loss_{j,q} = 1$ if quarter q reported earnings are negative, 0 otherwise.

[a] We report the mean coefficient estimates obtained from quarterly regressions (Q = 80) of each forecast measure (Panel A shows results for $Dispersion_{j,q}$, and Panel B shows results for $|FE|_{j,q}$) on the e-loadings (CAPM and 3-factor), *Age*, and *ln(Sales)*. We also report results that include the earnings surprise measures ($|UE^{SRW}_{j,q}|$, $NegUE^{SRW}_{j,q}$, and $Loss_{j,q}$) as independent variables. t-statistics are based on the standard errors of the 80 coefficient estimates.

must have data in event year $T = 0$. Therefore, the e-loading for year $T = +1$ is based on all firms that survived at least one year after listing. There is a clear downward trend in e-loadings as a function of *FirmAge* for both $e^{1f}_{j,T}$ and $e^{3f}_{j,T}$. To test the significance of this trend, we regress the average e-loading of firms of age τ ($\tau = 0,1,2,\ldots,39$) on *FirmAge*.[16] The results, also plotted in Figure 1, indicate a significant downward trend in e-loadings as a function of *FirmAge* (t-statistics, not reported, are −19.33 for $e^{1f}_{j,T}$ and −17.78 for $e^{3f}_{j,T}$). To control for over-time changes in the sample composition, we repeat our tests restricting the sample to firms that survived at least ten years following listing; results (not tabulated) show a significant decline in e-loadings for this sample. Overall, we interpret declines in

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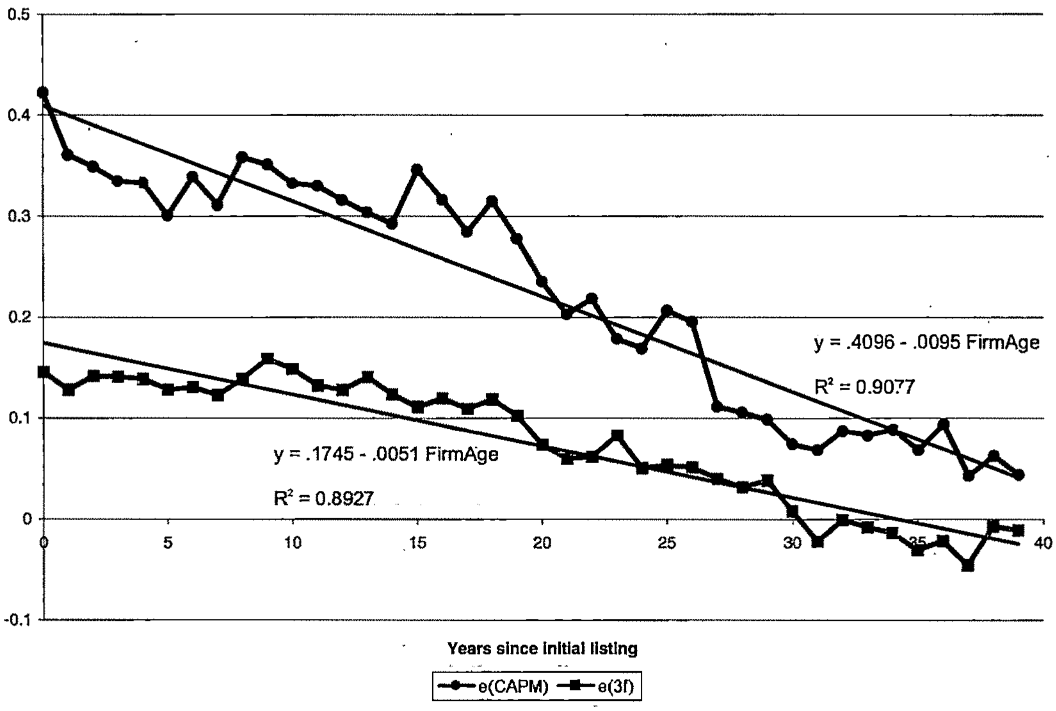
[16] We truncate observations at *FirmAge* = 39 because the sample size is small for values of *FirmAge* > 39; results are not sensitive to this truncation.

FIGURE 1
e-Loadings as Function of Firm Age



This shows the relation between firm age (measured as years since year 0, the year when the firm first listed on any of the U.S. stock exchanges) and the mean e-loading calculated across all firms in event year *T*. The graph demonstrates that e-loadings are highest in the early years of a firm's life and decline thereafter. The downward trends are reliably different from zero at the 0.0001 level.

e-loadings as a function of firm age as evidence that, as firm-specific information increases, investor uncertainty about reporting quality decreases.

Our second analysis investigates the prediction that e-loadings exhibit greater year-over-year stability for relatively mature firms than they do for relative young firms. In examining stability, we do not suggest that a firm's e-loadings *must* be the same from one year to the next; indeed, a benefit of e-loadings is that their estimation procedures do not force them to be constant over short intervals. As a practical matter, however, we would not expect to observe significant year-to-year changes in a firm's earnings quality (or investors' perceptions of that earnings quality) for firms with relatively stable reporting and information environments because, for such firms, relatively less new information about earnings quality is likely to be revealed each year. Our measure of the stability of e-loadings is the autocorrelation (AR1) parameter estimated for each cross-section of firms of *FirmAge* = 0,...,39 ($\omega_1^{\#}$):

$$e_{j,T}^{\#} = \omega_0^{\#} + \omega_1^{\#} e_{j,T-1}^{\#} + \zeta_{j,T}^{\#}. \qquad (8)$$

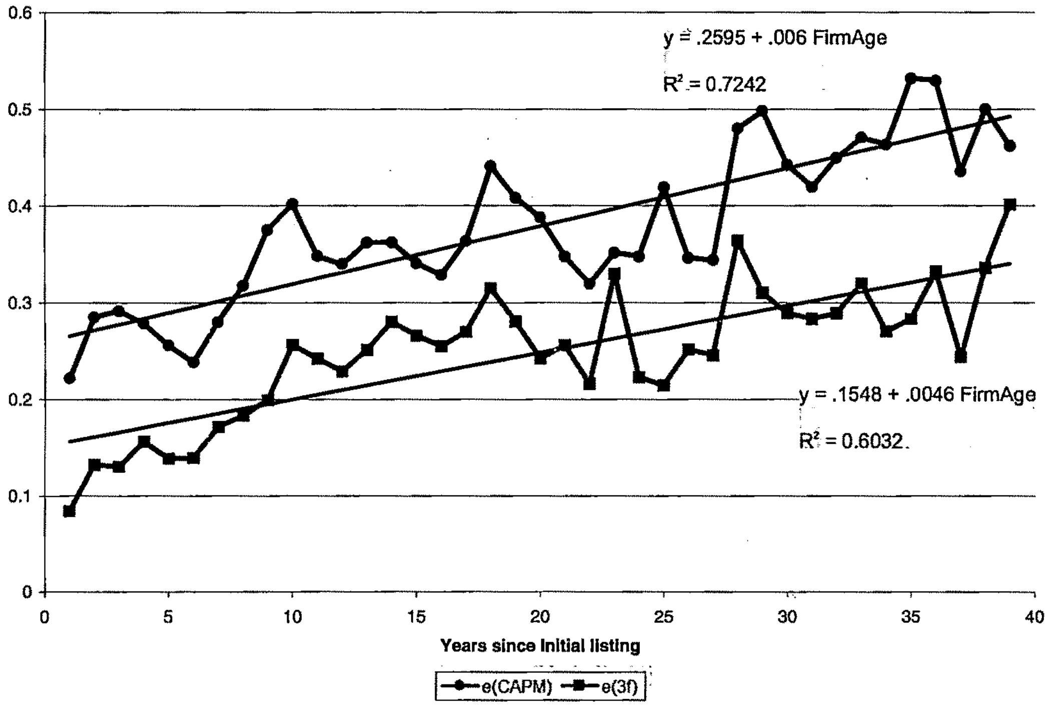
Figure 2 plots the estimates of $\omega_1^{\#}$ (representing each *FirmAge* cross-section) as a function of *FirmAge*; we also plot the results of regressing $\omega_1^{\#}$ on *FirmAge*. There is a pronounced increase in the stability of e-loadings: ω_1^{1f} and ω_1^{3f} increase from 0.22 and 0.08, respectively, for *FirmAge* = 0, to 0.46 and 0.40 for *FirmAge* = 39+ (t-statistics for these increases, not reported, are 9.86 and 7.98).

In summary, both the level of e-loadings and their stability exhibit systematic patterns related to firm age: (1) the level of e-loadings is highest at initial listing and declines as the firm matures; and (2) the year-over-year autocorrelation in e-loadings increases with firm age.

Patterns of e-Loadings Associated with Restatements, Lawsuits, and Bankruptcies

Our final analyses investigate whether e-loadings are higher in three situations characterized by objective evidence of poor financial reporting quality: financial statement restatements, shareholder lawsuits, and bankruptcies. These analyses are carried out in event time; so, for example, year 0 for the Restatement Sample (Lawsuit Sample) [Bankruptcy

FIGURE 2
Autoregressive Coefficient of e-Loadings as Function of Firm Age



This shows the relation between firm age and the stability of the year-over-year e-loading. Stability is measured as the AR1 parameter from a regression of $e^{\#}_{j,T}$ on $e^{\#}_{j,T-1}$, estimated in cross-section for all firms of *FirmAge* = 1,...,39. We report the AR1 parameter for each of the *FirmAge* groups. The graph demonstrates that the stability of e-loadings is lowest in early years of a firm's life and increase thereafter. The upward trends are reliably different from zero at the 0.0001 level.

Sample] is the year the restatement was announced (the lawsuit was filed) [the firm delisted due to bankruptcy]. Using an event-time analysis allows us to control for other (unrelated) influences on over-time patterns in e-loadings by comparing e-loadings of the event firms with e-loadings of non-event firms matched on calendar time. The event-time analysis also allows us to exploit the fact that each year there are many more firms that do not restate, are not sued, and do not go bankrupt than there are firms that experience any of these events. Specifically, for each of the Event Samples, we select 100 equal-size random samples of non-event firms, with a yearly distribution that is identical to that of the Event Sample. To the extent the smaller number of Event Sample observations is associated with greater measurement error in e-loadings than are the e-loadings of the larger set of non-event firms, equating the size of the Event and Non-Event Samples minimizes measurement error differences between the estimates derived from the two samples. Equating the yearly distributions also controls for calendar time influences affecting e-loadings. The 100 non-event samples create an empirical distribution of *non-event* e-loadings. A comparison of *event* e-loadings to this empirical distribution provides evidence on whether the e-loadings of the event firms are statistically distinguishable from those of firms that did not experience events associated with poor earnings quality.[17]

Our sampling and test procedures are the same for each of the three events, so we illustrate them only for the Restatement Sample. The Restatement Sample (described in more detail below) contains 788 firm-year restatements occurring between January 1, 1997 and June 30, 2002 with returns data. The yearly distribution of this sample is as follows: 1997 ($n = 82$), 1998 ($n = 87$), 1999 ($n = 151$), 2000 ($n = 162$), 2001 ($n = 199$), and the first half of 2002 ($n = 107$). We select 100 random samples of $n = 788$ total non-restating observations with the same yearly distribution (for example, 82 non-restating firms are randomly selected from 1997). Matching by year controls for events and conditions affecting all firms in a given year and provides the pseudo-event date (year 0) for the non-event firms. For each of the 788 observations in the Restatement Sample, we estimate Equations (2) and (3) for each year $T = -5,...,+1$, where $T = 0$ is the restatement year. We then calculate the mean value of the e-loadings across restatement firms, in event time, to obtain the mean value of $e^{1f}_{j,T}$ ($e^{3f}_{j,T}$) for each year $T = -5,...,+1$. For each of the 788 observations in each of the 100 Non-Restatement Samples, we perform the same estimation and averaging, yielding a series of 100 mean e-loadings for non-restating firms for each year $T = -5,...,+1$.

Our analyses of the Event Samples are intended to provide evidence on two issues: first, whether e-loadings increase for event firms relative to the non-event benchmark samples in the years preceding events that provide objective evidence of poor earnings quality ("pre-event tests"); and second, whether e-loadings decline in post-event periods for firms in the Restatement and Lawsuit Samples ("post-event tests"). (Post-event declines in e-loadings cannot be investigated for the Bankruptcy Sample because, by definition, these firms ceased trading, so returns data are not available to construct e-loadings in post-bankruptcy periods.) The pre-event and post-event tests are predicated on the assumptions that: (1) earnings quality deteriorated in the period preceding each of the three events and was, to some extent, corrected in the year after the event, and (2) investors were able to discern at least some portion of this deterioration. Our pre-event tests are clearly more sensitive to the second assumption than are the post-event tests, because the events we examine were all publicly disclosed.

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[17] Empirical distributions of comparison samples have been used in other contexts to establish a benchmark for assessing whether a test sample is distinctive. For example, Alford et al. (1993) use this approach to compare the value relevance and timeliness of earnings in the U.S. versus 16 non-U.S. jurisdictions. They compare measures of value relevance and timeliness for each of 16 non-U.S. test samples with an empirical distribution of 100 year-, size-, and industry-matched U.S. samples.

In terms of the first assumption, prior research provides mixed evidence on the association between earnings quality and events such as restatements, lawsuits, and bankruptcies. For the most part, this research examines changes in earnings quality metrics before and after restatements. For example, Anderson and Yohn (2003) compare bid-ask spreads and earnings response coefficients before and after restatements and find no change in spreads and a decline in ERCs. Moore and Pfeiffer (2004) find no significant difference in reporting aggressiveness (measured by abnormal accruals) before versus after a restatement. Taken together, these results suggest that correcting a GAAP violation by restating has either no effect, or a worsening effect, on earnings quality. These findings are at odds with intuition: assuming that GAAP earnings are of high quality, one would expect that when a GAAP violation is cured, the quality of the firm's financial reports improves not worsens.

In terms of the second assumption, prior research suggests that accounting data have predictive ability for the three events we study. Richardson et al. (2002) find that restatement firms have large accruals in the years of alleged manipulation, and that restatement firms try to maintain earnings patterns (such as consecutive positive earnings surprises and earnings growth). Research on lawsuit firms shows that firms sued over allegedly defective financial disclosures have systematic patterns in accounting variables, including higher levels of accounts receivable and inventory (Stice 1991), more income-increasing total accruals (Lys and Watts 1994), and more income-increasing abnormal accruals (Heninger 2001). Finally, numerous studies document systematic patterns in accounting data for bankrupt firms, including declines in profitability and increases in measures of financial distress (such as increases in book-to-market ratios and bankruptcy indices [Altman 1968, 1993]). Taken as a whole, we believe previous research supports the assumption that market participants have at least some information to form expectations about restatements, lawsuits, and bankruptcies before these events occur.

We analyze several predictions about the e-loadings of the Event and Non-Event Samples. First, we expect each Event Sample has larger e-loadings than its comparison Non-Event Samples in the event year, and in one or more years leading up to the event. Second, we expect that e-loadings of each Event Sample increase over time, relative to the e-loadings of comparison Non-Event Samples. For the Restatement and Litigation Samples we expect that e-loadings decline in year +1, consistent with year 0 actions correcting the financial reporting problems, e.g., restatement firms stop using incorrect GAAP and issue corrected financial statements. As noted earlier, we do not test whether e-loadings decline in year +1 for the Bankruptcy Sample because these firms cease trading after year 0.

Restatement Sample and Tests

Data on restatements come from the General Accounting Office's (GAO) study on Financial Statement Restatements, which covers January 1, 1997 through June 30, 2002. The GAO database identifies 919 restatement firm-years, representing 840 distinct firms. From this set, we identify all observations with usable identifier codes, resulting in a sample of 882 firm-year observations, for 794 firms. The distribution of this sample is shown in Panel A, Table 5. Imposing the requirement that firms have at least 100 daily returns in year T reduces the final Restatement Sample to 788 firm-year observations, representing 713 firms. Comparison of the yearly distributions of the two samples shown in Panel A indicates that the year-by-year distribution of this Restatement Sample is similar to the sample unconstrained by the requirement to have sufficient daily returns data.

TABLE 5
Yearly Distributions of Restatement, Lawsuit, and Bankruptcy Samples

Year	Original Sample	Firms with Returns Data	Year	Original Sample	Firms with Returns Data
Panel A: Restatement Sample					
1997	89	82	2000	190	162
1998	100	87	2001	216	199
1999	168	151	2002	119	107
			Total	882	788
Panel B: Lawsuit Sample					
1990	64	61	1997	72	68
1991	38	35	1998	99	93
1992	56	53	1999	83	74
1993	39	37	2000	80	74
1994	61	59	2001	68	64
1995	41	40	2002	108	94
1996	48	41	Total	857	793
Panel C: Bankruptcy Sample					
1970	6	4	1987	7	5
1971	4	1	1988	25	12
1972	2	1	1989	14	11
1973	8	6	1990	30	18
1974	6	4	1991	66	43
1975	6	4	1992	53	34
1976	1	0	1993	27	14
1977	4	4	1994	7	3
1978	8	3	1995	10	6
1979	3	1	1996	6	2
1980	5	3	1997	13	12
1981	9	5	1998	25	16
1982	16	13	1999	24	15
1983	17	6	2000	42	25
1984	12	8	2001	52	38
1985	11	6	2002	53	26
1986	15	10	2003	31	12
			Total	618	371

Sample descriptions: The Restatement Sample consists of firms with financial statement restatements over 1997–2002 (source: GAO [2002] study on financial restatements). The Lawsuit Sample consists of firms that experienced securities class action litigation over the period 1990–2002 where the lawsuit alleged one or more accounting violations. The Bankruptcy Sample consists of firms that delisted for bankruptcy reasons over the period 1970–2003. For each sample we report the number of firms in the original sample, and the number of firms with the necessary returns data (and so are available for our tests).

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Table 6, Panel A shows the e-loadings for the Restatement Sample and Non-Event Samples for event years −5 through +1. The e-loading for the Non-Event sample is the mean value of $e^{\#}_{j,T}$ calculated for the 100 matched samples. The column labeled "Diff." shows the difference between the mean value of $e^{\#}_{j,T}$ between the Event and Non-Event

TABLE 6

Comparison of e-Loadings of Event Firms (Restatement, Lawsuit, and Bankruptcy) with Non-Event Firms[a]

	Base Model: CAPM					Base Model: 3-Factor				
	Event Sample	Non-Event Sample	Diff.	t-stat	%tile	Event Sample	Non-Event Sample	Diff.	t-stat	%tile
Panel A: Restatement Sample										
Year −5	0.299	0.269	0.030	1.00	91	0.102	0.129	−0.027	−0.76	18
Year −4	0.318	0.288	0.030	1.08	94	0.112	0.143	−0.032	−0.98	10
Year −3	0.339	0.296	0.044	1.63	99	0.132	0.152	−0.020	−0.68	20
Year −2	0.352	0.277	0.074	2.89	100	0.159	0.164	−0.005	−0.17	42
Year −1	0.350	0.285	0.065	2.57	100	0.195	0.190	0.005	0.18	58
Year 0	0.390	0.268	0.122	4.32	100	0.315	0.207	0.108	3.57	100
Year +1	0.337	0.260	0.077	2.46	100	0.269	0.196	0.073	2.22	100
Panel B: Lawsuit Sample										
Year −5	0.276	0.297	−0.022	−0.59	20	0.028	0.128	−0.100	−2.61	1
Year −4	0.303	0.306	−0.003	−0.09	49	0.065	0.127	−0.061	−1.74	3
Year −3	0.312	0.297	0.015	0.49	76	0.031	0.127	−0.095	−2.98	1
Year −2	0.357	0.279	0.077	2.81	100	0.092	0.143	−0.052	−1.73	3
Year −1	0.421	0.273	0.148	5.70	100	0.157	0.155	0.002	0.08	60
Year 0	0.479	0.272	0.207	7.15	100	0.277	0.163	0.114	3.48	100
Year +1	0.474	0.252	0.222	6.56	100	0.286	0.152	0.134	3.74	100
Panel C: Bankruptcy Sample										
Year −5	0.488	0.332	0.156	3.81	100	0.243	0.113	0.129	2.86	99
Year −4	0.496	0.338	0.158	4.22	100	0.243	0.109	0.134	3.15	100
Year −3	0.512	0.321	0.191	5.24	100	0.275	0.118	0.157	3.90	100
Year −2	0.559	0.303	0.255	7.13	100	0.348	0.128	0.219	5.33	100
Year −1	0.619	0.295	0.324	7.75	100	0.516	0.136	0.379	7.74	100
Year 0	0.639	0.283	0.356	4.24	100	0.515	0.127	0.389	4.06	100

Sample descriptions: See Table 5 for descriptions of the three Event Samples (Restatement, Lawsuit, and Bankruptcy). For each Event Sample we randomly select 100 Non-Event Samples. Each of the 100 Non-Event Samples contains the same number of firms as the Event Sample; observations are selected from all non-event firms in year T so as to replicate the year-by-year distribution in the Event Sample.

[a] The columns labeled "Event Sample" show the mean e-loading for all Event firms in year $T = -5,\ldots,+1$ (for the Bankruptcy Sample, we end with year 0). The columns labeled "Non-Event Sample" show the average value of the e-loading calculated across the 100 Non-Event Samples. The columns labeled "Diff." show the difference between the mean values of the e-loadings for the Event and Non-Event Samples; the columns labeled "t-stat." show the t-statistic assessing the statistical significance of this difference (from zero). Finally, the columns labeled "%tile" show the percentile rank of the mean e-loading of the Event Sample within the empirical distribution of mean e-loadings for the Non-Event Samples. For example, a percentile rank of 100 implies that the Event Sample e-loading is the largest e-loading value of the distribution of mean e-loadings for the Non-Event Samples.

Samples. We evaluate the statistical significance of this difference in two ways. The first measure is a conventional t-statistic, based on the standard error of the mean e-loading of the Event Sample and the mean standard error of the 100 Non-Event Samples. The second is nonparametric and is reported in the column labeled "%tile": here we show the percentile rank of the Restatement Sample e-loading relative to the distribution of the 100 mean e-loading values of the Non-Event Samples. For example, a %tile value of 25 implies that the Restatement Sample e-loading ranks 25th from the bottom of the 100 mean Non-Event e-loadings.

Panel A, Table 6 shows that, as predicted, Restatement firms' e-loadings increase over the period preceding the restatement announcement year: the mean value of e^{1f} increases from 0.2994 in year -5 to 0.3899 in year 0, and declines to 0.3368 in year $+1$. For e^{3f} we observe a monotonically increasing trend, from 0.1018 to 0.3153 between years -5 to 0; the e-loading in year $+1$ declines to 0.2694. Over the same pre-event period, the results for e^{1f} for Non-Event firms show a slight decline, while results for e^{3f} show an increase up to year 0 (from 0.1292 in year -5 to 0.2072 in year 0), followed by a decline to 0.1963 in year $+1$. The trend in e^{3f} for the Non-Event Sample indicates a calendar-time effect that we control for by concentrating on the *differences* between the Event and Non-Event Samples.

Differences between e-loadings of the Event and Non-Event Samples are more pronounced over the entire period between years -5 and $+1$ for e^{1f} than for e^{3f}. For example, the value of e^{1f} for the Restatement Sample is at or above the 91st percentile of the Non-Event Samples in every year considered, and parametric tests yield t-statistics well above 2.0 in years -2 through $+1$. In contrast, values of e^{3f} for the Restatement Sample are below the median of the Non-Event Samples in years -5 through -2 and t-statistics exceed 2.0 only in years 0 and $+1$. Clearly, however, differences are statistically reliable for both e^{1f} and e^{3f} in years 0 and $+1$, with the largest differences observed in year 0. The significant difference in year $+1$ suggests that market participants continue to perceive restatement firms as having poorer earnings quality in the year after the restatement occurs.

Lawsuit Sample and Tests

Data on class action securities litigation are from Woodruff-Sawyer & Co. and from the Securities Class Action Clearinghouse. In total, there are 2,707 firm-year lawsuit observations over the period January 1, 1990 through April 24, 2003. Because we are interested in lawsuits over financial reporting issues, we focus on the subset of 1,067 lawsuits alleging accounting violations or concerns. We find usable identifier data for 857 lawsuits; the requirement of 100 daily returns in year T reduces the final Lawsuit Sample to 793 lawsuits against 751 distinct firms. Table 5, Panel B shows the distribution of the Lawsuit observations across the sample years, before and after imposing the returns requirement; the year-by-year distributions of the two samples are similar.

Examination of the e-loadings of the Lawsuit Sample (Table 6, Panel B) suggests an increase in e^{1f} from 0.2756 to 0.4792 for lawsuit firms over years -5 to year 0, and a (slight) decline in year $+1$. For the matched Non-Event firms, there is a near-monotonic (but limited) decline in e^{1f} over years -5 to $+1$. The difference in e^{1f} between the Event and Non-Event firms shows a pronounced upward trend, from -0.0215 (year $+5$) to 0.2070 (year 0), further increasing to 0.2217 in year $+1$. The three-factor e-loadings increase for both Lawsuit firms and Non-Event firms over the period -5 to 0; however, the increase is much more pronounced for Lawsuit firms. Comparing the two samples, the Lawsuit firms exhibit lower values of e^{3f} than Non-Event firms until year -1, and Lawsuit firms' e^{3f} are significantly larger than Non-Event firms e^{3f} in years 0 and $+1$. Nonparametric tests place the Lawsuit firms' one-factor (three-factor) e-loadings at the 100th percentile of the Non-Event Samples in years −2 through +1 (years 0 and +1). The weight of the evidence supports the view that e-loadings do not differ between Lawsuit and Non-Event firms in most pre-event years, but differ dramatically in years 0 and +1 when Lawsuit firms' e-loadings exceed those of Non-Event firms.

Bankruptcy Sample and Tests

We identify 618 firms that went bankrupt during 1970–2003 using delisting codes obtained from the CRSP database. Requiring returns data reduces the sample to 371 bankruptcy observations.[18] The yearly distributions of both samples are similar (Panel C, Table 5). Results comparing the e-loadings of Bankrupt firms with Non-Event firms are shown in Panel C, Table 6. We observe an upward trend in both e^{1f} and e^{3f} for Bankrupt firms over the event years, with the highest e-loadings found in the year before and the year of bankruptcy. Non-Event firms show a downward trend in e^{1f} over this same time frame, and a modest upward trend in e^{3f}. The difference in e^{1f} between Event and Non-Event firms increases monotonically from 0.1558 in year −5 to 0.3558 in year 0 while the difference in e^{3f} also increases monotonically, from 0.1293 in year −5 to 0.3888 in year 0. Both the parametric t-test and the nonparametric test indicate that the average e-loading is substantially higher for the Bankruptcy Sample than for the Non-Event Samples in every year examined.

Summary of Results of Event Tests

The weight of the evidence for the three Event Samples supports the following inferences. First, e-loadings are significantly larger for Event firms than for Non-Event firms in the event year (for all Event Samples) and in the year following (for the Lawsuit and Restatement Samples). Second, for all Event Samples there is an increase in the mean e-loading of the Event firms over years −5 to 0. For the Restatement Sample, we also observe the predicted decline in e-loading between year 0 and year +1; however, this pattern is not systematically observed for the Lawsuit Sample. Third, the difference between the e-loadings of the Event firms and the Non-Event firms tends to increase over years −5 to year 0 for all Event Samples. This pattern, which controls for changes in e-loadings due to economy-wide factors, indicates that investors perceive declining earnings quality for Event firms over the period leading up to and culminating in the event year. This pattern is most pronounced, and begins earliest, for the Bankruptcy Sample where, arguably, investors have access to relatively more pre-event information.[19]

As a check on our treatment of the events as independent, we calculated the overlap in firms represented in the three Event Samples. The Bankruptcy Sample displays minimal overlap with either the Restatement Sample (4.3 percent) or the Lawsuit Sample (7.8 percent). For the Restatement and Lawsuit Samples, the overlap is 25 and 27 percent, respectively. We address this sample dependence by repeating the tests in Table 6 excluding all firms that appear in more than one Event Sample. The results (not reported) are similar to those for the full sample, in terms of both the magnitudes of the coefficients and their statistical significance.

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[18] The decline in sample size is due to the fact that requiring at least 100 daily returns in calendar year T effectively eliminates firms that went bankrupt in the first five months of year T.

[19] Note that the results for e^{3f} control for changes in other risk factors over the event periods. The importance of this is best demonstrated by the bankruptcy sample, where firms' risk profiles change as they approach bankruptcy. Consistent with Fama and French's (1995) finding that h-loadings proxy for relative distress, we find that h-loadings increase for firms approaching bankruptcy.

VI. EXTENSIONS

We conduct several extensions of our main analyses. First, to examine the possibility that our results are driven by the subset of firms with data on *AQ* (i.e., the AQ Sample), we repeat our tests on the subset of firms in the Returns Sample but not in the AQ Sample. Results (not tabulated) are similar in all respects to those reported. This finding demonstrates that the significant expansion of sample size made possible by not requiring data on *AQ* produces e-loadings that are as reliable as e-loadings generated for the AQ Sample itself.

Second, we consider how well e-loadings for other earnings quality proxies perform, relative to e-loadings based on *AQfactor*. Our aim is to assess whether *any* returns representation of earnings quality performs equally well along the dimensions we examine. We calculate mimicking factors for the six earnings quality proxies considered in Section V, yielding a returns representation for each: *Factor(k)*, k = *Persistence*, *Predictability*, *Smoothness*, *Value Relevance*, *Timeliness*, and *Conservatism*. We substitute *Factor(k)* for *AQfactor* in Equations (2) and (3), generate firm-specific e-loadings for each of the k factors, $e(k)^{\#}_{j,T}$, and then repeat our tests k times, substituting $e(k)^{\#}_{j,T}$ for $e^{\#}_{j,T}$. The results (not tabulated) show that e-loadings based on other measures of earnings quality generally show similar patterns as e-loadings based on *AQfactor*, except that e-loadings for *Timeliness* and *Conservatism* show weaker effects. Of the remaining e-loadings, those for *Smoothness* and *Persistence* are most similar, in magnitude and statistical significance, to those documented for *AQfactor*. (This result is consistent, in hindsight, with *Smoothness* and *Persistence* being closer to accruals quality in capturing the precision of earnings with respect to cash flows). There is no evidence that any of the k e-loadings systematically dominates the *AQ*-based e-loading ($e^{\#}_{j,T}$).

Third, we consider an abnormal accruals metric, equal to the absolute value of the performance-matched abnormal accrual ($|AA_{j,T}|$) estimated in annual cross-sections for each of the 48 Fama-French (1997) industries containing at least 20 firms. We calculate a mimicking factor based on $|AA_{j,T}|$, $Factor(|AA|)$, and repeat all tests on the resulting e-loadings based on this factor, $e(|AA|)^{\#}_{j,T}$. The results show that while $e(|AA|)^{\#}_{j,T}$ shows similar patterns as $e^{\#}_{j,T}$, the results are generally weaker for this measure. Based on this evidence, as well as that documented for $e(k)^{\#}_{j,T}$, we conclude that e-loadings based on accruals quality performs at least as well as, or better than, e-loadings based on other measures of earnings quality.

Fourth, we consider whether shorter estimation periods produce e-loadings that reliably capture earnings quality. Shorter estimation intervals are advantageous because they place fewer restrictions on the sample and allow for greater time-variation in e-loadings; they are disadvantageous because they increase the noise in the resulting e-loadings. Our extension investigates how well e-loadings perform when they are calculated at the quarterly level; here, we require a firm to have at least 45 daily returns in calendar quarter Q to estimate its quarterly e-loading, $e^{\#}_{j,Q}$. Results (not tabulated) show that $e^{\#}_{j,Q}$ exhibit similar patterns as found for $e^{\#}_{j,T}$. Further, while the confidence intervals are wider, all of the results for $e^{\#}_{j,Q}$ are reliably significant at the 10 percent level or better. We conclude that returns-based representations of earnings quality calculated using as few as 45 daily returns reliably capture aspects of earnings quality.

Our fifth and final extension considers the overlap between properties of e-loadings and properties of broader measures of information risk or information asymmetry. The two most prominent measures are bid-ask spreads and probability of informed trading (PIN) scores. The properties of both measures have been explored primarily in the market microstructure literature (see Huang and Stoll [1997] for bid-ask spreads and Easley, Kiefer, and O'Hara

[1996] for PIN scores). Huang and Stoll show that the bid-ask spread has three components: (1) processing costs, (2) inventory costs, and (3) costs due to adverse selection/information asymmetry. Based on extensive data analysis (they analyze all trades during one year for their sample firms), they conclude that the average adverse information component is 9.6 percent of the spread. This finding is based on very large firms; it is conceivable that the information asymmetry component is relatively more important for stocks of smaller and younger firms (Easley, Kiefer, O'Hara, and Paperman 1996). Probability of informed trading (PIN) scores incorporate information on spreads as well as order flow information to generate an estimate of the probability that a given trade order originates from a privately informed trader. Firms with larger PIN scores are viewed as having more information asymmetry among investors than firms with smaller PIN scores.

We use bid ask-spread data for Nasdaq firms for 1983–2003,[20] and we use PIN scores for NYSE/AMEX firms for the period 1983–2001.[21] As an initial assessment of the overlap between e-loadings and both bid-ask spreads and PIN scores, we correlate firm-year e-loadings with firm-year average bid-ask spreads and firm-year PIN scores. To control for scale effects, we divide the bid-ask spread by the midpoint of the bid and ask. The Pearson (Spearman) correlation between e-loadings and bid-ask spreads is 0.14 (0.12); the correlation between e-loadings and PIN scores is similar: 0.13 (0.16).[22] All correlations are significant at the 0.01 level. We interpret these results as indicating that while there is statistically nonzero overlap between e-loadings and these measures of information asymmetry, its economic magnitude is limited. This result is consistent with Huang and Stoll's (1997) conclusion that a small portion of the bid-ask spread (about 10 percent) can be linked to information asymmetry.

We probe these results further by orthogonalizing e-loadings with respect to bid-ask spreads and PIN scores, respectively. Our goal is to assess the sensitivity of the main results to the shared variation between e-loadings and bid-ask spreads, and between e-loadings and PIN scores. Specifically, we run the following regressions by year:

$$e^{3f}_{j,T} = \lambda^{BA}_{0,T} + \lambda^{BA}_{1,T} BA_{j,T} + \varepsilon^{BA}_{j,T} \tag{9}$$

$$e^{3f}_{j,T} = \lambda^{PIN}_{0,T} + \lambda^{PIN}_{1,T} PIN_{j,T} + \varepsilon^{PIN}_{j,T} \tag{10}$$

where BA stands for the bid-ask spread and PIN denotes the PIN score. The residual from Regression (9) or (10) is the portion of the e-loading that is orthogonalized with respect to bid-ask spreads or PIN scores, $\hat{\varepsilon}^{Orthog}_{j,T}$. Note that this procedure is conservative in the sense that any shared variation between e-loadings and bid-ask spreads (or between e-loadings and PIN scores) will be attributed to the information asymmetry measure and thus "removed" from the e-loading.[23] Note further that due to data availability, $\hat{\varepsilon}^{Orthog}_{j,T} = \hat{\varepsilon}^{BA}_{j,T}$ exists for Nasdaq firms for 1983–2003, while $\hat{e}^{Orthog}_{j,T} = \hat{e}^{PIN}_{j,T}$ exists for NYSE/AMEX firms for 1983–2001. For these subperiods, we repeat our main tests setting $\hat{e}^{Orthog}_{j,T} = \hat{e}^{BA}_{j,T}$ for Nasdaq firms and $\hat{e}^{Orthog}_{j,T} = \hat{e}^{PIN}_{j,T}$ for NYSE/AMEX firms. (Results are similar if we restrict the sample to only bid-ask spread orthogonalized firms or to PIN-score orthogonalized firms.) Table 7 contains the results from tests based on $\hat{e}^{Orthog}_{j,T}$. For brevity, we tabulate one market participant test (analyst forecast dispersion), one test of a situation associated with objectively poor earnings quality (restatements), and one over-time test (firm age); all conclusions generalize to other tests. As is evident from Table 7, the behavior of $\hat{e}^{Orthog}_{j,T}$ is very similar to the behavior of the non-orthogonalized e-loading, $e^{\#}_{j,T}$, documented in our main tests.

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[20] CRSP reports bids and asks for all securities listed on the Nasdaq National Market since November 1982, and all Nasdaq securities since June 15, 1992. The inside quotation (the highest bid and lowest ask) is used as the closing bid and ask. The source data for February 1986 is missing; hence, this month is also missing on CRSP.

[21] These data are from Søren Hvidkjær's website: http://www.smith.umd.edu/faculty/hvidkjaer/. To be included in year T, a stock must further have at least 60 days with quotes or trades in that year. The final sample has between 1,863 and 2,414 stocks per year.

[22] We cannot assess the overlap between bid-ask spreads and PIN scores because the former are available for Nasdaq firms and the latter for NYSE/AMEX firms. However, for a sample of firms where both measures are available, Hasbrouck (2004) reports correlations between closing spreads and PIN scores of 0.40 (Pearson) and 0.66 (Spearman).

[23] We also performed tests that orthogonalized e-loadings with respect to loadings on factor-mimicking portfolios based on bid-ask spreads and PIN scores. Results (not reported) are similar in all respects to those where the orthogonalization is performed with respect to the construct itself (BA or PIN).

It is possible that the effects of bid-ask spreads and PIN scores are better captured by

TABLE 7

Tests of e-Loadings Orthogonalized with Respect to Spreads and PIN Scores[a]

Panel A: Regression of Forecast Dispersion on e^{Orthog} and Control Variables

	Base Model: 3-Factor			
Variable	**Coefficient**	**t-statistic**	**Coefficient**	**t-statistic**
e^{Orthog}	0.0014	10.15	0.0007	6.21
Age	0.0000	−4.68	0.0000	−4.21
ln(size)	−0.0002	−7.98	0.0000	−1.38
AbsUE	NA	NA	0.0457	8.79
NegUE	NA	NA	0.0008	10.66
Loss	NA	NA	0.0035	17.01
Adj. R^2	0.0342		0.2817	

Panel B: Comparison of e^{Orthog} of Restatement Firms with Non-Event Firms

	Base Model: 3-Factor				
	Event Sample	**Non-Event Sample**	**Diff.**	**t-stat.**	**%-tile**
Year −5	0.0065	−0.0043	0.0107	0.2825	61
Year −4	0.0058	−0.0027	0.0085	0.2461	66
Year −3	0.0155	−0.0024	0.0178	0.5978	76
Year −2	0.0307	−0.0035	0.0343	1.2275	97
Year −1	0.0292	−0.0021	0.0313	1.1491	89
Year 0	0.1137	−0.0027	0.1163	3.6674	100
Year +1	0.1012	0.0005	0.1007	2.5855	100

Panel C: Regression of e^{Orthog} on Firm Age

	Coefficient	**t-statistic**
Intercept	0.0365	5.58
Age	−0.0031	−10.89
Adj. R^2	0.7509	

Sample description and variable definition: The sample consists of all firm-year observations with data on $\hat{e}^{Orthog}_{j,T}$ and the other variables in each test, over 1983–2003. $\hat{e}^{Orthog}_{j,T}$ is firm j's e-loading in year T, orthogonalized with respect to bid-ask spreads or PIN scores.

[a] Panel A shows the results of regressing analyst forecast dispersion on $\hat{e}^{Orthog}_{j,T}$ as well as control variables. Panel B shows values of $\hat{e}^{Orthog}_{j,T}$ for event and non-event firms in the five years leading up to, the year of, and the year after restatement announcements. Panel C reports the results of regressing $\hat{e}^{Orthog}_{j,T}$ on firm age.

loadings on factor-mimicking portfolios for each construct rather than on the constructs themselves. Using returns representations of these constructs also allows us to extend the sample to all firms, not just the sample of firms for which ask spreads (BA) and PIN scores are available; further, the resulting BA-loadings and PIN-loadings are directly comparable to e-loadings. To form the BA-mimicking portfolio, we begin by calculating, for each firm-month, the average bid-ask spread; using the monthly average removes some of the variation in daily bid-ask spreads, stabilizing the portfolio. BA-loadings ($BA^{\#}_{j,T}$) are then calculated using the same procedures as described for the e-loadings (i.e., same ranking procedures, returns measurement with a three-month lag, and taking long-short positions). Because PIN-scores are available at the firm-year level, we follow the same procedure as *AQ* for creating a mimicking factor and estimating firm-year specific PIN-loadings ($PIN^{\#}_{j,T}$). Measures of both $BA^{\#}_{j,T}$ and $PIN^{\#}_{j,T}$ are available for the same set of firms for which we have e-loadings: 1983–2003 for $BA^{\#}_{j,T}$ and 1983–2001 for $PIN^{\#}_{j,T}$. We then repeat our tests replacing $e^{\#}_{j,T}$ with $BA^{\#}_{j,T}$ (or $PIN^{\#}_{j,T}$). In all cases, the results (not tabulated) show weaker patterns, and weaker statistical significance of those patterns, relative to tests based on $e^{\#}_{j,T}$.[24] Based on these results, as well as those in Table 7, we conclude that the earnings quality effects captured by e-loadings are not driven by information asymmetry effects captured by bid-ask spreads and PIN scores.

VII. SUMMARY AND CONCLUSION

We describe a returns-based representation of earnings quality, in the form of the coefficient estimate (the e-loading) from firm-specific regressions of daily excess returns on a factor-mimicking portfolio capturing earnings quality, controlling for other risk factors. Our analysis is predicated on Francis et al.'s (2005) analysis of accruals quality as a valid empirical measure of information risk as a priced factor. Theoretical support for information risk as a priced factor is provided by analytical models, for example, Easley and O'Hara (2004), Leuz and Verrecchia (2005), and Lambert et al. (2005). Each of these studies posits a different information risk pricing mechanism: Easley and O'Hara (2004) provide a trading model in which better quality reporting reduces the information risk faced by investors who have access to public signals only; Leuz and Verrecchia (2005) provide a real-effects model in which higher quality reporting supports a better alignment between investors and managers with respect to investment decisions (a solution to an agency problem); and Lambert et al. (2005) posit a framework in which information risk may be priced because of the inability to fully specify a forward-looking CAPM beta. We do not distinguish among various reasons put forward by analytical research for *why* information risk is priced (although we regard this as an interesting research question). Moreover, our reading of these papers suggests that different reasons for why information risk is priced may coexist; for example, Lambert et al. (2005) explicitly note that apart from their own model, effects such as those in Easley and O'Hara (2004) can influence firms' costs of capital.

Our tests build on prior empirical research demonstrating that information risk (as proxied by accruals quality) is priced; we extend this work by developing and validating a returns-based representation of information risk that has several empirical advantages relative to accounting-based representations. While our main tests focus on accruals quality, we note that accruals quality is not the *only* valid empirical measure of information risk;

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[24] Further, in the market participant tests where it is straightforward to include $e^{\#}_{j,T}$ as well as $BA^{\#}_{j,T}$ (or $PIN^{\#}_{j,T}$) as independent variables, we find that $e^{\#}_{j,T}$ is always more important than either of the information asymmetry measures in explaining market reactions to unexpected earnings conveyed by earnings announcements, analyst forecast dispersion, and analyst forecast accuracy.

moreover, accruals quality may capture effects other than information risk (although we attempt to control for known effects in our tests). As a returns-based representation of earnings quality, e-loadings are subject to similar caveats.

We document that e-loadings are a reliable returns-based representation of earnings quality as measured by accruals quality. That is, e-loadings are positively associated with other measures of earnings quality; they proxy for the uncertainty in earnings as viewed by investors and by analysts; and they exhibit expected over-time patterns as a function of firm age. Further, in settings where earnings quality has arguably changed (restatements, lawsuits, and bankruptcies), e-loadings show predictable patterns both over-time and in relation to e-loadings for firms which did not experience these events. We find similar patterns for e-loadings based on both annual and quarterly estimations, and for e-loadings that are value-weighted and orthogonalized with respect to firm size versus equal-weighted and not size-orthogonalized.

The demonstrated reliability of yearly and quarterly e-loadings has at least two implications for future research. First, e-loadings impose fewer sampling restrictions, relative to measures of earnings quality calculated using a time-series of either accounting data (such as persistence, predictability, and smoothness) or a combination of accounting and market data (such as value relevance, timeliness, and conservatism). The longer the time-series required to obtain a reliable estimate of these other measures, the fewer firms with the required series (leading to smaller samples), the greater is the survivorship bias (leading to samples skewed toward larger, successful firms), and the more static (that is, the less time-specific) is the resulting measure. In contrast, because reliable e-loadings can be calculated using only returns data, the resulting samples are more representative of the population of firms (that is, the samples can include smaller, younger firms that lack a time-series of Compustat data) and the e-loadings themselves are not mechanically constrained to be constant or slow to change. The greater representativeness of the sample affords both greater power (due to more variation in attributes of sample firms) and greater generalizability. In addition, because e-loadings can be estimated reliably over intervals as short as a quarter, they can be used to examine the effects of events (such as the promulgation of new accounting standards) or actions (such as earnings management) that might be expected to influence earnings quality at points in time or over fairly short intervals.

A second implication derives from the significant associations between e-loadings and other measures of earnings quality, such as persistence, smoothness, predictability and, to a lesser extent, value relevance, timeliness and conservatism. To the extent these other measures are viewed as capturing aspects of a firm's "true" earnings quality, positive associations between them and e-loadings suggest that investors' *perceptions* of earnings quality are rational. Stated differently, if investors were unable to discern (or unable to discern correctly) the pricing implications of earnings quality, we would not expect e-loadings to correlate with non-perception-based measures of earnings quality or with innate factors that have been shown to explain such measures.

We believe that returns-based measures of earnings quality are most appropriate in research settings that analyze changes in financial reporting quality, either because of a required change in reporting or because of a voluntary reporting or disclosure decision. Returns-based measures are likely to be less appropriate, or inapplicable, in research settings that are characterized by: shifts in both information risk and fundamental risks (such as mergers); explicit predictions about signs and/or magnitudes of effects (there is no theory that predicts the signs or magnitudes of e-loadings); or a requirement to separate total earnings quality into discretionary and innate portions. We also acknowledge that our results do not speak to whether e-loadings based on accruals quality are the preeminent returns-based measure of earnings quality. It seems unlikely that any single measure would be best in all research settings and, in fact, our results suggest that e-loadings based on other measures of earnings quality, such as persistence and smoothness, also exhibit substantial reliability. However, our results also indicate that e-loadings based on accruals quality perform at least as well as other measures in the contexts that we examine.

REFERENCES

Alford, A., J. Jones, R. Leftwich, and M. Zmijewski. 1993. The relative informativeness of accounting disclosures in different countries. *Journal of Accounting Research* 31 (Supplement): 183–223.

Altman, E. 1968. Financial ratios, discriminant analysis and the prediction of corporate bankruptcy. *Journal of Finance* 23: 589–609.

———. 1993. *Corporate Financial Distress and Bankruptcy.* 2nd edition. New York, NY: John Wiley & Sons.

Anderson, K., and T. Yohn. 2003. The effect of 10-K restatements on firm value, information asymmetries, and investors' reliance on earnings. Working paper, Georgetown University.

Basu, S. 1997. The conservatism principle and the asymmetric timeliness of earnings. *Journal of Accounting and Economics* 24 (1): 3–37.

Botosan, C., and M. Plumlee. 2005. Assessing alternative proxies for the expected risk premium. *The Accounting Review* 80 (1): 21–53.

Bowen, R., A. Davis, and D. Matsumoto. 2002. Do conference calls affect analysts' forecasts? *The Accounting Review* 77 (2): 285–316.

Carhart, M. 1997. On persistence in mutual fund performance. *Journal of Finance* 52: 57–82.

Chen, S., T. Shevlin, and Y. Tong. 2004. What is the information content of dividend changes? A new investigation of an old puzzle. Working paper, University of Washington.

Daniel, K., and S. Titman. 1997. Evidence on the characteristics of cross sectional variation in stock returns. *Journal of Finance* 52: 1–33.

Dechow, P., and I. Dichev. 2002. The quality of accruals and earnings: The role of accrual estimation errors. *The Accounting Review* 77 (Supplement): 35–59.

Easley, D., N. Kiefer, and M. O'Hara. 1996. Cream-skimming or profit-sharing? The curious role of purchased order flow. *Journal of Finance* 51: 811–833.

———, ———, ———, and J. Paperman. 1996. Liquidity, information, and less-frequently traded stocks. *Journal of Finance* 51: 1405–1436.

——— and M. O'Hara. 2004. Information and the cost of capital. *Journal of Finance* LIX (4): 1553–1583.

Fama, E., and J. MacBeth. 1973. Risk, return and equilibrium: Empirical tests. *Journal of Political Economy* 81: 607–636.

———, and K. French. 1993. Common risk factors in the returns on stocks and bonds. *Journal of Financial Economics* 33: 3–56.

———, and ———. 1995. Size and book-to-market in earnings and returns. *Journal of Finance* 50: 131–155.

———, and ———. 1997. Industry costs of equity. *Journal of Financial Economics* 43: 153–193.

———, and ———. 2004. New lists: Fundamentals and survival rates. *Journal of Financial Economics* 73: 229–269.

Francis, J., R. LaFond, P. Olsson, and K. Schipper. 2004. Costs of equity and earnings attributes. *The Accounting Review* 79 (4): 967–1010.

———, ———, ———, and ———. 2005. The market pricing of accruals quality. *Journal of Accounting and Economics* 39 (2): 295–327.

Hasbrouck, J. 2004. Trading costs and returns for US equities: The evidence from daily data. Working paper, New York University.

Heflin, F., K. R. Subramanyam, and Y. Zhang. 2003. Regulation FD and the financial information environment: Early evidence. *The Accounting Review* 78 (1): 1–37.

Heninger, W. 2001. The association between auditor litigation and abnormal accruals. *The Accounting Review* 76 (1): 111–126.

Hribar, P., and D. Collins. 2002. Errors in estimating accruals: Implications for empirical research. *Journal of Accounting Research* 40: 105–134.

Huang, R., and H. Stoll. 1997. The components of the bid-ask spread: A general approach. *The Review of Financial Studies* 10: 995–1034.

Imhoff, E., and G. Lobo. 1992. The effect of ex ante earnings uncertainty on earnings response coefficients. *The Accounting Review* 67: 427–439.

Lambert, R., C. Leuz, and R. Verrecchia. 2005. Accounting information, disclosure and the cost of capital. Working paper, University of Pennsylvania.

Lamont, O., C. Polk, and J. Saá-Requejo. 2001. Financial constraints and stock returns. *The Review of Financial Studies* 14: 529–554.

Lang, M. 1991. Time-varying stock price response to earnings induced by uncertainty about the time-series properties of earnings. *Journal of Accounting Research* 29 (2): 229–257.

———, and R. Lundholm. 1996. Corporate disclosure policy and analyst behavior. *The Accounting Review* 71 (4): 467–492.

Leuz, C., and R. Verrecchia. 2005. Firms' capital allocation choices, information quality, and the cost of capital. Working paper, University of Pennsylvania.

Lewellen, J. 1999. The time-series relations among expected return, risk and book-to-market. *Journal of Financial Economics* 54: 5–43.

Lys, T., and R. Watts. 1994. Lawsuits against auditors. *Journal of Accounting Research* 32 (Supplement): 65–93.

McNichols, M. 2002. Discussion of: The quality of accruals and earnings: The role of accrual estimation errors. *The Accounting Review* 77 (Supplement): 61–69.

Moore, E., and R. Pfeiffer. 2004. The effects of financial statement restatements on firms' financial reporting strategies. Working paper, University of Massachusetts.

Newey, W., and K. West. 1987. A simple, positive semi-definite, heteroscedasticity and autocorrelation consistent covariance matrix. *Econometrica* 55: 703–708.

Pastor, L., and P. Veronesi. 2003. Stock valuation and learning about profitability. *Journal of Finance* 58: 1749–1789.

Richardson, S., I. Tuna, and M. Wu. 2002. Predicting earnings management: The case of earnings restatements. Working paper, University of Pennsylvania and Hong Kong University of Science and Technology.

Stice, J. 1991. Using financial and market information to identify pre-engagement factors associated with lawsuits against auditors. *The Accounting Review* 66 (3): 516–553.

Subramanyam, K. R. 1996. The pricing of discretionary accruals. *Journal of Accounting and Economics* 22: 249–281.

Xie, H. 2001. The mispricing of abnormal accruals. *The Accounting Review* 76: 357–373.

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Does Hiring a New CFO Change Things? An Investigation of Changes in Discretionary Accruals

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ABSTRACT: The recent spate of fraudulent financial reporting in the U.S. has drawn attention to the fact that the Chief Financial Officer (CFO) has a substantial amount of control over a firm's reported financial results. This paper examines the changes in discretionary accruals surrounding the appointment of a new CFO. Using a sample of 712 companies that appointed a new CFO in the period 1994 to 2000, we find that discretionary accruals decreased significantly following the appointment of a new CFO. Our tests indicate that this reduction is significantly greater for our group of CFO-hiring firms than for a control group of non-hiring firms, and that the changes are not driven by a concurrent appointment of a new Chief Executive Officer (CEO). We also find that our results are largely driven by firms that hire a new CFO from outside the company. Our study extends earlier research by providing empirical evidence that individuals in CFO positions wield significant influence over the firm's reported financial results and that a firm's discretionary accruals are significantly reduced surrounding the appointment of a new CFO.

Keywords: *CFO turnover; discretionary accruals; corporate governance.*

Data Availability: *The data used in this study are publicly available from the sources indicated in the text.*

I. INTRODUCTION

Congressional and media attention has recently been directed toward Chief Financial Officers (CFOs) and the role they play in creating and disseminating accurate or fraudulent financial information. While inaccurate information is not a new problem for the financial markets, the recent level of Congressional, media, and public outcry over financial reporting misbehavior by U.S. corporate officials is largely unprecedented. Fueled by the financial frauds and subsequent bankruptcy filings of several high-profile corporations, legislators and the business press have examined the financial reporting of these

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companies and have begun to lay blame for the scandalous reporting not only with the Chief Executive Officers (CEOs) who have final oversight authority for the company, but also with the CFOs. What was uncovered during these corporate reporting investigations has revealed for some, and reinforced for others, the significant role the CFO plays in contemporary financial reporting.

These investigatory findings also fueled the Congressional resolve to adopt legislation to hold not only the CEO, but also the CFO personally accountable for the accuracy and completeness of the financial information released by the company (Geiger and Taylor 2003). The eventual outcome of those Congressional hearings was a formal requirement in the Sarbanes-Oxley Act of 2002 that CEOs *and* CFOs *personally* certify to the material accuracy and completeness of the financial information and disclosures released by the company. Thus, CFOs have now been legislatively elevated to the same level of financial oversight responsibility as was formerly viewed as the responsibility of the CEO. Personal certification by the CEO and CFO for the financial information released by their corporations is also consistent with, and reflective of, Congress's view that specific individuals at these high-level executive positions have significant influence over the publicly disclosed information of their companies.

Despite a longstanding acknowledgement regarding the involvement of CFOs in the financial reporting process, little academic research has been conducted to examine the effect of the CFO on a corporation's reported financial results. There has been considerable research regarding the effect of new CEOs on corporate financial reporting (Strong and Meyer 1987; DeAngelo 1988; Pourciau 1993; Murphy and Zimmerman 1993) and the relationship between the timing of management actions and reported financial results (Aboody and Kasznik 2000; Bartov and Mohanran 2004; Yermack 1997), yet there has been very little investigation regarding the relationship between new CFOs and reported financial information (Mian 2001).

As implied by the new certification requirements, if individuals appointed as CFOs are in a position to significantly effect the reporting of their company's financial condition, then a change in CFO personnel may well lead to different financial results to be reported under similar business conditions. Accordingly, in this study we investigate the relationship between appointing a new CFO and the changes in a company's reported discretionary accounting accruals surrounding this turnover event. If individual CFOs maintain an ability to effectively shape the financial information reported, then the exercise of this influence would be manifest in changes in a company's reported discretionary accounting accruals surrounding the change in personnel. Thus, an examination of changes in discretionary accruals around the hiring of a new CFO lends empirical evidence to the ability of the individual CFOs to significantly effect their firms' reported financial statements. Motivation for our study comes not only from the paucity of research in this area, but also from the heightened and ongoing interest of U.S. regulators, legislators, and investors regarding corporate financial reporting and the involvement of the CFO for the release of accurate financial information (SEC 2000, 2002; U.S. Senate 2002; U.S. House of Representatives 2002; Geiger and Taylor 2003).

Evidence from a sample of 712 companies that appointed a new CFO in the years 1994–2000 indicates that discretionary accruals (calculated using the cross-sectional Jones [1991] model), decreased significantly after the appointment of a new CFO compared to other non-hiring firms. Specifically, we find that our sample of hiring firms report significantly higher discretionary accruals in the year prior to the new CFO appointment and significantly lower discretionary accruals in the first full reporting year under the new CFO. We test separately the change in accruals surrounding the CFO turnover event and find that there are significantly greater reductions for our CFO hiring sample compared to our sample of non-hiring firms during the same periods.

In multivariate tests we control for factors found by prior research to be associated with levels of discretionary accruals and reach similar conclusions regarding the reduction in discretionary accruals for our CFO-hiring firms. We then assess whether our aggregate findings are driven by the concurrent appointment of a new CEO in our examination period. Our results using only the 390 firms that did not have concurrent CEO turnover in our CFO turnover window produce similar results and provide strong evidence that our findings are not driven by the concurrent hiring of a new CEO.

We then assess whether our results are different for firms appointing CFOs from within or hiring them from outside the company. Our results indicate that hiring an external CFO is associated with significant reductions in discretionary accruals, while appointing an internal CFO is not. We also find that companies hiring a new CFO directly from their own external CPA firm exhibit similar reductions in discretionary accruals as do other firms appointing a new CFO.

Next we conduct several robustness and additional tests. To ensure that our results are not driven by the performance characteristics of our sample firms, we examine changes in performance-adjusted discretionary accruals metrics (Kothari et al. 2005) for our sample of hiring firms and for all non-hiring firms with available data, and obtain consistent results. We then consider whether our results are driven by the firm size, or the year of CFO appointment (early or late in our sample time period). After including control variables in our regressions for these potential confounds, we obtain consistent and significant results for our CFO hire firms. Finally, we consider whether our results regarding the reduction of abnormal accruals are uniform across the three reporting-year period surrounding the CFO appointment. Additional subperiod regressions present evidence that the reductions found in accruals are fairly evenly distributed across our multiyear event window. In sum, these additional analyses indicate that our overall findings are robust across differing discretionary accruals metrics, variable specifications and across various segments of our sample of CFO-hiring firms.

Our study contributes to the corporate governance and discretionary accounting accruals literatures by assessing whether the appointment of a new individual to the executive position of CFO is associated with significant changes in the company's reported discretionary accruals. In contrast to the extant literature, we provide broad evidence regarding the significance of appointing a new CFO on a company's reported financial statements. Our findings present consistent support for the hypothesis that individual CFOs exhibit a considerable amount of influence over their firm's reported financial information, and that after the appointment of a new individual as CFO, companies report significantly lower discretionary accounting accruals.

The remainder of the paper is organized as follows. Section II discusses the related literature and develops our hypotheses. Section III presents our research method and results. Section IV presents our additional analyses and robustness tests. The final section discusses the research limitations and presents the conclusions.

II. PRIOR LITERATURE AND HYPOTHESIS DEVELOPMENT

Prior Literature

The practice of corporate accounting and reporting has received considerable attention in the academic literature (Bartov et al. 2002; Matsumoto 2002; Pincus and Rajgopal 2002) and in the business press (Byrnes et al. 1998; Collingwood 2001; Smith 2004). Investors, regulators, audit committees, and external auditors are all concerned with individuals and companies attempting to manage earnings by manipulating levels of reported accounting accruals (Bartov and Mohanran 2004; Dechow et al. 1996; Levitt 1998, 2000). This concern has spawned a considerable amount of research into the quality of corporate earnings and the existence of earnings management through the use of accounting accruals (cf., Aboody and Kasznik 2000; Ashbaugh et al. 2003; Barton and Simko 2002; Dechow and Dichev 2002). Prior research has examined the relationship between financial statement manipulations and various company characteristics (DeFond and Jiambalvo 1994; Dechow et al. 1996; Teoh et al. 1998), audit firm characteristics (Francis et al. 1999; Bartov et al. 2000; Chung and Kallapur 2003), boards of directors and audit committee characteristics (Beasley et al. 1999; Carcello and Neal 2000; Klein 2002; Xie et al. 2003), and CEO actions and characteristics (Strong and Meyer 1987; DeAngelo 1988; Pourciau 1993; Murphy and Zimmerman 1993; Bartov and Mohanran 2004).

Of particular import to our study is the extant research on the appointment of a new CEO and changes in reported financial results immediately afterward. This body of research has consistently found that new CEOs are more likely to reduce reported income results in their initial year—the "big bath"—than companies not hiring a new CEO (Strong and Meyer 1987; DeAngelo 1988; Pourciau 1993; Murphy and Zimmerman 1993; Denis and Denis 1995). However, while this occurrence is well documented for CEO changes, and despite their dominant and significant financial oversight role, there has been very little direct investigation of the CFO in the financial reporting process or on levels of reported accounting accruals surrounding to the appointment of a new CFO.

In the first broad examination of CFO appointments, Mian (2001) examined CFO hiring company characteristics and where the new CFOs came from and where the former CFOs went in the corporate executive turnover process. He concluded that CFO turnover is often punitive in that it is regularly preceded by poor stock price and operating income performance, and is also preceded by high CEO turnover. However, his study focused on the reason for CFO turnover and where these individuals came from and went to, and did not assess the effect of the turnover process on the company's subsequent financial results.

More recent studies in accounting have begun to assess the impact of hiring a new CFO on levels of accounting accruals in isolated hiring situations. Menon and Williams (2004), Dowdell and Krishnan (2004), and Geiger et al. (2005) examine the impact of a company hiring an individual directly from their external audit firm—i.e., the hiring practice referred to as the "revolving door" (Clikeman 1998)—which is now restricted under the Sarbanes-Oxley Act. The Menon and Williams (2004) study examines companies that have appointed former partners of their external audit firm to corporate officer positions like CFO, Controller, etc., as well as to board of directors' positions. The focus of their research is on the effect of former audit-firm-related partners on a company's reported financial results and, accordingly, do not restrict their study solely to companies hiring directly from their audit firm. However, they do test a subsample of companies appointing individuals from their CPA firm into CFO and similar financial reporting oversight positions and find that discretionary accounting accruals are significantly higher for these firms than for their control samples.[1] Unfortunately, even this subsample includes all former partners of the company's external audit firm regardless of length of employment at the company or whether the individual came directly from the CPA firm, effectively precluding a direct test of the "revolving door" hires on reported accounting accruals.

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[1] Additionally, Lennox (2005) examines audit firm decision making involving companies hiring former audit firm colleagues. His findings indicate that auditors are less likely to issue a going-concern modified audit report to companies who have an executive who is a former employee of the audit firm, consistent with the argument of impaired audit quality in these situations.

The Dowdell and Krishnan (2004) and Geiger et al. (2005) studies provide a more direct assessment of the "revolving door." Dowdell and Krishnan (2004) examine appointments to CFO positions, and include individuals at all ranks at their former audit firm; Geiger et al. (2005) examine only former managers and partners of CPA firms, but include appointments to other financial reporting positions such as Controller, VP-Finance, etc. Dowdell and Krishnan (2004) conclude that levels of discretionary accruals are higher for former audit personnel that recently joined their publicly traded clients. In contrast, Geiger et al. (2005) find no differences between their experimental sample of revolving-door hiring firms and control samples of non-revolving-door hiring firms with respect to changes in absolute accruals or in the nonoperating accruals measure proposed by Givoly and Hayn (2000). Thus, these recent studies have taken different approaches (e.g., sample selection procedures and accruals metrics) to examine financial reporting surrounding this now restricted hiring practice, and have reached differing conclusions. To date, however, no study has presented a targeted examination of the broad relationship between CFO appointments and a company's reported financial results surrounding these hirings as part of the executive turnover process.

Changes in Discretionary Accruals and CFO Turnover

The association between appointing a new CFO and changes in the company's financial reporting are assessed in our study by examining abnormal, or discretionary, accounting accruals surrounding the appointment of a new individual. The academic literature has differentiated between nondiscretionary (or normal) and discretionary (or abnormal) components of accounting accruals. Nondiscretionary accruals are the expected, or normal, level of accruals for the company based on factors such as type of operating industry, company size, and revenue growth (Jones 1991; DeFond and Jiambalvo 1994; Kothari et al. 2005). The discretionary, or abnormal, accruals for a company are the unexpected component reported by the company and are the difference between actual levels reported and the level expected to be reported by the company for the period. Previous researchers have argued that the level of these discretionary accruals reported by the company is a reflection of management's use of the "financial reporting discretion" inherent in GAAP to either increase or decrease reported net income (Schipper 1989; Jones 1991; Levitt 2000; DeFond and Park 2001). Since discretionary accruals are believed to be subject to more reporting judgment and are, therefore, more easily manipulated than are nondiscretionary accruals, if individual CFOs, by virtue of their oversight position alone, can significantly impact the reported financial results of their firms, then we would expect that influence to be manifest in changes in the levels of discretionary accruals reported by companies surrounding a change in personnel.

Consistent with DeFond and Jiambalvo (1994), Becker et al. (1998), and Chung and Kallapur (2003), we examine signed current discretionary accruals. These earlier researchers have argued that companies are rarely sued for booking accruals to reduce earnings, so examining signed discretionary accruals is the most appropriate measure of the intentional influence of management on financial reporting. Further, Becker et al. (1998), DeFond and Park (2001), and Ashbaugh et al. (2003) also note that current discretionary accruals should be the focus of accruals research in that current accruals are where management has the most immediate reporting discretion to manipulate reported financial results. Therefore, if individuals intend to actively manage the firm's reported results, then the effect would be most salient in the changes in levels of signed current discretionary accruals. This is especially germane to our study in that we assess corporate turnover at the CFO position and examine a relatively short period of years surrounding the appointment of a new CFO.

Examining signed current discretionary accruals is also consistent with the evidence presented in Kreutzfeldt and Wallace (1986), Kinney and Martin (1994), and Nelson et al. (2002), which indicate a majority of a client company's "judgmental errors" identified by their external auditors, or where auditors perceived there to be manipulation attempts by clients, are located in the areas of revenue recognition and current asset and liability accounts. The findings of these audit adjustment studies lend credence to the argument that corporate managers, in general, more readily use current accounts in order to try to manipulate the company's reported results. Accordingly, we examine the effect of a change in CFO on levels and changes in levels of signed current discretionary accruals surrounding the turnover event.

As in DeFond and Jiambalvo (1994), in order to assess the influence of individual CFOs over corporate financial reporting, we present data on the year-end current discretionary accruals positions of our hiring firms for periods immediately before and after appointing a new CFO. Additionally, similar to Geiger et al. (2005), we also examine the changes in levels of these reported accruals from the period immediately prior to the turnover event (i.e., $t-1$) to the first full reporting period after the new CFO (i.e., $t+1$) by using the change in levels of signed current discretionary accruals (i.e., $t+1$ minus $t-1$) as our change metric. Our multiyear approach and the specific examination of changes in discretionary accruals enables us to provide a better assessment of the effect individuals have on reported accruals than examining accruals levels at the end of a single year (e.g., time t). Further, since the new CFO might take over at any time throughout year t, it also becomes difficult to accurately assess the ability of the new CFO to impact the reported results in their initial year of hire, so examining discretionary accruals levels at time t alone is problematic. A more robust assessment of the influence of individuals on reported discretionary accruals is to examine not only the levels of discretionary accruals, but also the changes for the period immediately surrounding CFO turnover (the period $t-1$ to $t+1$).

Hypotheses

If an individual in the position of CFO is able to exert significant influence over the reported financial results of their company, as discussed previously and as presumed in the new certification requirements in the Sarbanes-Oxley Act, then we would expect a reduction in abnormal accruals over the period $t-1$ to $t+1$ for two reasons. First, the former CFO has incentives to increase accruals in $t-1$ and, second, the new CFO has incentives to reduce accruals in $t+1$.

As found in prior research on audit adjustments (Kinney and Martin 1994; Nelson et al. 2002) an incumbent CFO, on average, has incentives to increase discretionary accruals in order to present a more favorable picture to the markets and to maximize compensation in the form of bonuses and pay increases tied to achieving financial targets (Healy and Wahlen 1999). Bonuses and compensation contracts for company executives are often predicated on the prior-year's financial or stock price performances and afford the incumbent CFO significant incentive to reach corporate profit and financial performance targets in order to maximize his/her personal wealth. Additionally, leaving or retiring CFOs have similar and often heightened incentives to maximize corporate profits and their salaries in the years immediately preceding their retirement or departure (Burgstahler and Dichev 1997; Cheng 2004). This is due to the fact that pension and severance contracts are often based in large part on salary and compensation levels in the few years immediately preceding retirement or separation (Bartov et al. 2002; Dechow and Sloan 1991). These incentives for wealth maximization would result in higher than expected positive discretionary accruals from incumbent CFOs.

An additional incentive for CFOs to increase discretionary accruals levels is to try to maintain their employment with the firm. Mian (2001), in his study of CFO turnover, concludes that CFO turnover is often punitive—resulting from poor prior reported financial statement income levels or below average stock price performance. If firms replace CFOs for reporting poor financial or stock price performance, then CFOs have a tremendous amount of incentive to use discretionary accruals to project the most favorable presentation of their company to the financial and equity markets in order to try and maintain their CFO positions. Thus, we believe there is considerable incentive for incumbent CFOs to report positive financial results and to use discretionary accruals to aid in that attempt.

Conversely, new CFOs may have incentives to reduce discretionary accruals. The extant research on the appointment of a new CEO and changes in reported financial results clearly demonstrates that companies appointing new CEOs are more likely to reduce reported income in their initial years of the CEO—the "big bath"—than similar companies not hiring a new CEO (Strong and Meyer 1987; Pourciau 1993; Murphy and Zimmerman 1993; Denis and Denis 1995). Reporting low income levels in the early years of employment with a new company enables executives to show more substantial financial reporting improvements over their tenure than if the initial financial results were stronger. We suspect that CFOs would maintain similar reporting incentives as CEOs in this respect and tend to reduce reported levels of income upon their arrival at their new firm in order to show more positive financial results in the periods under their tenure. Accordingly, newly appointed CFOs would also have incentive to initially reduce discretionary accruals and adopt a conservative reporting posture immediately after they assume reporting responsibility in order to lay claim to more improved financial results under their tutelage.

We expect that the combined effect of these incentives would be that companies hiring new CFOs would report significant reductions in the level of discretionary accruals in the periods surrounding the turnover event. Thus, the first hypothesis examined in our study is:

> **H1:** Companies that appoint a new CFO report significant reductions in discretionary accruals compared to other non-hiring firms.

Alternatively, since they are not directly involved in profit generating activities, and in contrast to the Congressional presumption in Sarbanes-Oxley Act, if CFOs are not able to exert influence over the reported financial results of their company, then we would expect to find no significant change in levels of discretionary accruals surrounding the appointment of a new individual to the position of CFO.

Our second hypothesis addresses differences in discretionary accruals changes resulting from new CFOs coming from different sources (Mian 2001). Parrino (1997) finds evidence that the origin of the new CEO (e.g., hired from outside or appointed from within) is associated with differences in company financial performance after the new CEO takes over. Specifically, he finds significantly greater changes in reported financial performance immediately after a new CEO takes over who was hired externally as compared to an internal CFO appointment. Accordingly, we examine if similar hiring source differences may exist with CFO appointments.

In the examination of changes in discretionary accruals surrounding a CFO appointment, we would also expect that promoting or appointing an individual from within the company would result in smaller changes in financial reporting than hiring a new CFO from outside the company. Individuals joining the company for the first time would be expected to bring significantly different perspectives to the financial reporting issues facing the company than those held by current employees. These greater differences in financial reporting perspective would result in larger modifications of existing policies, and cause significantly greater changes in reported accruals after the new external individual takes over as compared to appointing an internal person, who may be more wedded to the company's existing reporting positions. Thus, our second hypothesis:

> **H2:** Companies hiring external individuals as new CFOs exhibit greater reductions in discretionary accruals compared to companies appointing internal individuals.

A related issue to this second hypothesis is whether individuals hired directly from the company's external audit firm (i.e., the "revolving door" hires) report similar changes in discretionary accruals compared to other CFO hires. Accordingly, we provide additional analyses that examine our sample of CFO hire firms for evidence of a "revolving door" effect on changes in reported discretionary accounting accruals.

III. METHOD AND RESULTS

Sample

To identify newly appointed CFOs we conducted a full search of the Dow Jones Interactive database for the keywords "CFO" or "Chief Financial Officer" in the period 1994–2000. To ensure that we could obtain the required financial data to calculate the discretionary accruals estimation models, we confined our searches to NYSE-, AMEX-, and NASDAQ-listed firms. After identifying those releases that include new CFO appointments, we then searched background biographies (in Factiva) to identify the former employers of the newly hired individuals. Our initial search identified 1,305 companies that announced appointments of individuals to CFO positions during 1994–2000, of which 1,143 companies were included in Compustat. Consistent with prior discretionary accruals studies, due to idiosyncratic and industry-specific financial reporting issues, we excluded financial services and utilities industry firms (SIC codes 6000–6999 and 4900–4999, respectively). As in Mian (2001), and required by our second hypothesis, our test sample of companies hiring a new CFO consists of only those cases where the individual's background prior to appointment was unambiguous. Additionally, because we need to calculate discretionary accruals models for three consecutive annual reporting years, firms were excluded if they did not have at least four consecutive years of financial data available in Compustat, leaving 712 companies identified as having appointed a new CFO with all the necessary data (hereafter, our HIRE sample).

Table 1 presents the distribution of our sample of our HIRE companies (i.e., companies appointing a new CFO) across years as well as industries. As presented in Panel A, our HIRE sample is slightly skewed toward the more recent years in our assessment period. Panel B indicates that our sample is fairly representative of the distribution of firms across industries represented in the Compustat database. As noted in Panel B, after eliminating financial services and regulated industries, the average distribution of firms in the Compustat database from 1994 to 2000 was: SIC 10–19: 6.1 percent; SIC 20–39: 45.9 percent; SIC 40–48: 10.8 percent; SIC 50–59: 11.4 percent; and SIC 70–89: 25.8 percent. Our sample distribution was: SIC 10–19: 3.9 percent; SIC 20–39: 47.9 percent; SIC 40–48: 7.0 percent;

TABLE 1
Hire Distribution

Panel A: Hire Distribution by Event-Year

Year	Observations
1994	48 (7.0%)
1995	62 (8.7%)
1996	87 (12.2%)
1997	100 (14.0%)
1998	116 (16.3%)
1999	159 (22.3%)
2000	139 (19.5%)
Total	712 (100.0%)

Panel B: Hire Distribution by Two-Digit SIC Industry

SIC Codes	Industry	HIRE Sample Observations	Compustat Proportions
10–19	Mining and Construction	28 (3.9%)	6.1%
20–39	Manufacturing	341 (47.9%)	45.9%
40–48	Transportation	50 (7.0%)	10.8%
50–59	Wholesale and Retail	100 (14.1%)	11.4%
70–89	Services	193 (27.1%)	25.8%
10–89	Total	712 (100.0%)	100.0%

Distribution for sample of 712 companies that appointed a new CFO in the period 1994–2000.

SIC 50–59: 14.1 percent; and SIC 70–89: 27.1 percent. Thus, we find no strong evidence of disproportionate industry clustering in our sample of firms.

Our sample of non-hiring control firms includes all firms in Compustat with necessary data to calculate the discretionary accruals models and all control variables used in our multivariate regressions. Additionally, all financial services and utilities industry firms were excluded (SIC codes 6000–6999 and 4900–4999, respectively). Based on this search procedure we were able to identify 31,325, 32,445, and 32,932 firms in $t-1$, t, and $t+1$ that did not disclose the appointment of a new CFO that were captured in the Dow Jones Interactive database. These non-HIRE firms serve as our population of yearly control firm samples.[2] Table 2 presents descriptive data for our sample of HIRE and non-HIRE companies for each of the three years.[3]

Mean and median comparisons indicate that HIRE companies are generally larger in terms of market value of equity, have lower sales growth and are more likely to have obtained additional financing or entered into an acquisition compared to non-HIRE companies. As indicated in Table 2, we find no significant differences between our HIRE and

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[2] While we used the Dow Jones Interactive database to identify our HIRE firms, undoubtedly there are companies that did not publicly disclose the hiring of CFOs, or where the release was not included in the Dow Jones Interactive database. However, inclusion of these unidentified hiring firms in our control sample of non-HIRE firms would bias our results against finding a HIRE firm effect.

[3] We winsorize outliers (top and bottom 1 percent in each year) for discretionary accruals, size, book-to-market ratio, distress (measured using Zmijewski's [1984] model), cash flow scaled by total assets, and sales growth percentage. Our results are substantively unchanged if we delete these firm outliers in selecting our samples.

TABLE 2
Descriptive Statistics

	Time ($t-1$)		Time (t)		Time ($t+1$)	
	Mean	**Median**	**Mean**	**Median**	**Mean**	**Median**
HIRE Firms (n = 712)						
LN Market value equity	5.111***	4.861***	5.065***	4.849***	4.918***	4.839***
Book-to-Market Equity	0.487	0.417	0.561**	0.441	0.494	0.484**
Current ratio	2.718	2.001	2.612***	1.896	2.524***	1.801**
Debt/assets	0.229	0.186	0.245	0.199	0.257	0.197
ROA	−0.056**	0.022	−0.071***	0.018	−0.076***	0.009
CFFO/Total Assets	0.003	0.047*	−0.002*	0.052	0.012***	0.052
Sales Growth	0.297	0.133*	0.243***	0.112	0.154***	0.063***
Financing Dummy	0.498***	0.000***	0.471	0.000	0.401**	0.000**
Acquisition Dummy	0.345***	0.000***	0.337**	0.000***	0.310	0.000
All Non-HIRE Firms (n = 31,325/32,445/32,928)						
LN Market value equity	4.531	4.447	4.517	4.440	4.501	4.445
Book-to-Market Equity	0.486	0.434	0.495	0.439	0.463	0.439
Current ratio	2.832	1.979	2.873	1.974	2.296	1.966
Debt/assets	0.242	0.188	0.248	0.187	0.261	0.184
ROA	−0.081	0.031	−0.115	0.025	−0.144	0.018
CFFO/Total Assets	−0.004	0.062	−0.020	0.056	−0.031	0.053
Sales Growth	0.294	0.111	0.330	0.111	0.323	0.099
Financing Dummy	0.446	0.000	0.451	0.000	0.447	0.000
Acquisition Dummy	0.282	0.000	0.290	0.000	0.292	0.000

*, **, *** Denotes significance at the 10%, 5%, and 1% levels, respectively.
p-values were calculated using two-tailed t-tests (Wilcoxon rank sum tests) for differences in means (medians).
Time (t) is the year of the new CFO appointment for the sample of 712 HIRE firms.

non-HIRE firms in terms of median return on assets (ROA); however, we find that our HIRE firms have a higher mean ROA than the non-HIRE firms. This pattern is indicative of our sample of HIRE firms having fewer extreme negative accruals than are present in the non-HIRE firms.[4] In order to control for these financial performance differences, we include an ROA variable in our multivariate regressions. Thus, we report the results of using the modified cross-sectional Jones (1991) model, discussed in the next section.[5]

Univariate Tests—H1

Consistent with prior research, we use the cross-sectional version of the Jones (1991) model introduced by DeFond and Jiambalvo (1994) to calculate discretionary current accruals. Predicted accruals for our study are estimated cross-sectionally using the following model:

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[4] Deleting firms with extreme ROA measures (i.e., below 1 percent and above 99 percent) results in mean ROAs that are not significantly different ($p > .10$) between the samples. Rerunning our analyses on these reduced samples produces essentially the same results as those reported.

[5] As noted in our robustness tests presented in a subsequent section, we obtain similar results when using the ROA performance-adjusting procedure advocated in Kothari et al. (2005).

$$AC_{i,t}/TA_{i,t-1} = \alpha_i[1/TA_{i,t-1}] + \beta_{1i}[(CREV_{i,t} - CAR_{i,t})/T_{i,t-1}] + e_i \tag{1}$$

where *AC* is our accruals measure defined as change in noncash current assets minus the change in current liabilities excluding the current portion of long-term debt (all from the operations section of the statement of cash flow); *TA* is total assets; *CREV* is change in revenue; and *CAR* is change in accounts receivable, for firm *i* at year *t*. The model is estimated cross-sectionally for every two-digit SIC classification using all available firms except the sample firm. The estimated coefficients from the industry regression are used to estimate predicted accruals for the sample. Discretionary accruals (DA) are actual accruals minus predicted accruals.[6] Table 3 presents data for DA for our samples across the three years surrounding a change in CFO.

Comparing the means and medians in Panel A of Table 3 for the period $t-1$ we find that our sample of 712 CFO HIRE firms have significantly higher DA (mean of 3.01 percent of total assets) compared to the DA for our sample of non-HIRE firms (mean of 1.41 percent of total assets).[7] This difference in DA between these two groups (mean and median) is significant at the $p < .01$ level. Thus, companies hiring a CFO exhibit significantly higher levels of discretionary accruals for the last reporting year under the former CFO (i.e., at time $t-1$) and present evidence that discretionary accruals are used to increase income for the HIRE firms in the year prior to a CFO change. This finding is consistent with CFOs attempting to increase discretionary accruals to either maximize current or retirement compensation levels, or to maintain their position within the firm, or both.

In time *t* we find that our HIRE firms report a mean DA of 1.75 percent of total assets, compared to the 3.01 percent reported in the year before, and compared to the 1.33 percent level reported by the sample of non-HIRE firms in time *t*. This reduction in mean DA for the HIRE firms is significant at the $p < .01$ level, yet the change in DA from $t-1$ to *t* for the non-HIRE firms is not significant ($p > .25$). The difference in DA between the HIRE and non-HIRE firms is also not significant in time *t* ($p > .10$). The time *t* results indicate that the levels of discretionary accruals were substantively reduced for the HIRE firms but not for the non-HIRE firms, and that the significant differences in DA between our two samples at time $t-1$ disappear in this subsequent year.

The time $t+1$ results reported in Table 3 indicate that the HIRE firms, on average, report even further reduced levels of DA (mean of 0.01 percent of total assets) compared to their previously reported DA levels in $t-1$ and *t* ($p < .01$). As expected, in time $t+1$ the HIRE firms reported significantly lower levels of DA compared to the non-HIRE firms ($p < .01$). Thus, while the non-HIRE firms report relatively stable levels of DA over our three year examination period (1.41 percent, 1.33 percent and 0.98 percent for $t-1$, *t*, and $t+1$, respectively) the HIRE firms consistently and significantly reduced levels of DA over the period surrounding a new CFO appointment (from 3.01 percent to 1.75 percent to 0.01 percent for $t-1$, *t*, and $t+1$, respectively).

Panel B of Table 3 presents our direct tests of the changes in DA (ΔDA) over time for our HIRE and non-HIRE firms. The mean (median) change of -2.98 percent (-1.72 percent) for the HIRE firms across our time frame is significantly different from zero ($p < .01$), indicating that HIRE firms report significantly reduced levels of signed DA. This

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[6] All firms with available data are used to calculate DA, even if they did not have all data available on the control variables to enable them to be used in the later analyses. Industries (based on two-digit SIC codes) with fewer than ten firms were deleted.

[7] If we restrict our control sample of non-HIRE companies to only those 25,399 that had all data needed to calculate DA for all three years, then we obtain substantially the same results as presented.

TABLE 3
Descriptive Information on Discretionary Accruals

Panel A: Discretionary Accruals (DA) by Time[a]

	Time ($t-1$)		**Time (t)**		**Time ($t+1$)**	
	Mean	**Median**	**Mean**	**Median**	**Mean**	**Median**
All Non-HIRE Firms (n = 31,325/32,445/32,928)						
DA	0.0141	0.0090	0.0133	0.0088	0.0098	0.0072
All HIRE Firms (n = 712)						
DA	0.0301***	0.0168***	0.0175	0.0146*	0.0001***	0.0023***
HIRE Firms without CEO turnover in *t*−1, *t*, or *t*+1 (n = 390)						
DA	0.0357***	0.0177***	0.0236**	0.0147**	−0.0006**	0.0011**
HIRE Firms with CEO turnover in *t*−1, *t*, or *t*+1 (n = 322)						
DA	0.0232	0.0153	0.0101	0.0146	0.0009	0.0029
HIRE Firms with External Hires (n = 574)						
DA	0.0331***	0.0196***	0.0187	0.0167**	−0.0017***	0.0020***
HIRE Firms with Internal Hires (n = 138)						
DA	0.0175	0.0059	0.0133	0.0088	0.0080	0.0029

Panel B: Changes in Discretionary Accruals (ΔDA): Time ($t+1$) minus Time ($t-1$)[b]

	Mean	**Median**
All Non-HIRE Firms (n = 25,399)		
ΔDA	−0.0086	−0.0025
All HIRE Firms (n = 712)		
ΔDA	−0.0298***	−0.0172***
HIRE Firms without CEO turnover in *t*−1, *t*, or *t*+1 (n = 390)		
ΔDA	−0.0360***	−0.0214***
HIRE Firms with CEO turnover in *t*−1, *t*, or *t*+1 (n = 322)		
ΔDA	−0.0222	−0.0142*
HIRE Firms with External Hires (n = 574)		
ΔDA	−0.0347***	−0.0221***
HIRE Firms with Internal Hires (n = 138)		
ΔDA	−0.0094	−0.0029

***, **, * Denotes significant difference compared to the sample of Non-HIRE firms at the 1%, 5% and 10% levels, respectively.

p-values were calculated using two-tailed t-tests (Wilcoxon rank sum tests) for differences in means (medians).

[a] Discretionary accruals by time generated by the Jones (1991) model; DA is abnormal accruals.

[b] Changes in discretionary accruals from time ($t-1$) to time ($t+1$). Time (t) is the year of the new CFO appointment for the sample of CFO Hire firms and non-HIRE firms.

is in contrast to the 25,399 non-HIRE firms that have all necessary data to calculate accruals for all three years that show a mean (median) change of only −0.86 percent (−0.25), which is not different from zero ($p > .25$), and is significantly smaller than the reduction in DA exhibited by the HIRE firms ($p < .01$). These results, on average, amount to an additional reduction of $20.25 million in DA for our HIRE firms compared to our non-HIRE firms over the three-year reporting period surrounding a CFO change.

These univariate results on the changes in DA provide consistent support that our sample of non-HIRE firms maintained the same level of discretionary accruals over our examination period; however our HIRE firms significantly reduced their discretionary accruals surrounding the appointment of a new CFO. Next, we conduct several multivariate analyses to further investigate our univariate results.[8]

Multivariate Tests—H1

To provide additional examination of the impact of hiring a new CFO on changes in a firm's reported discretionary accruals, we use the following multivariate model to control for factors found in prior research to be related to levels of discretionary accruals:

$$\begin{aligned} DA = {} & a + b_1 HIRE + b_2 MVE + b_3 BM + b_4 DISTRESS + b_5 CFFO \\ & + b_6 GROWTH + b_7 FINANCE + b_8 ACQ + b_9 ROA \end{aligned} \qquad (2)$$

where:

DA = discretionary accruals estimated from the cross-sectional Jones model;
HIRE = 1 if company hired a CFO, 0 otherwise;

and the control variables are

MVE = log of the market value of equity;
BM = book-to-market equity ratio;
DISTRESS = financial distress measure (calculated from Zmijewski 1984);
CFFO = cash flow from operations divided by total assets;
GROWTH = sales growth rate;
FINANCE = 1 if number of o/s shares increased by at least 10 percent or long-term debt increased by at least 20 percent during the year;
ACQ = 1 if the company engaged in an acquisition; and
ROA = return on assets from the prior year.

In our model we include controls for company size, financial condition, operating cash flow, sales growth, and structural changes due to significant new financing or acquisitions. While our accruals measures are scaled for size, it is still possible that scaled accruals are related to firm size. We use the log of market value of equity (*MVE*) at the end of the period as our measure of firm size. *BM* is the proportion of book value to market value and represents the growth opportunities available to the firm. Prior research has found size and *BM* to be negatively related to discretionary accruals (Ashbaugh et al. 2003; Butler et al. 2004; Menon and Williams 2004).

We include a measure of financial condition (*DISTRESS*) because of concerns that the Jones (1991) model may overestimate accruals for poorly performing companies (Dechow et al. 1995; Kothari et al. 2005). We use Zmijewski's (1984) measure of financial distress as our firm condition metric. Greater values of *DISTRESS* indicate higher levels of financial stress present in the firm. Reynolds and Francis (2000), Ashbaugh et al. (2003), and Menon and Williams (2004) find evidence that financial health is negatively associated with discretionary accruals. *CFFO* is our measure of operating cash flow scaled by total assets.

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[8] The sections in Table 3 examining differences in CEO/non-CEO hiring samples and internal/external hires are presented here and discussed in subsequent sections of the paper.

Prior research has shown that *CFFO* is negatively related to abnormal accruals (Ashbaugh et al. 2003; Becker et al. 1998; Chung and Kallapur 2003; Frankel et al. 2002). *GROWTH* is the percentage growth in sales over the period. *GROWTH* has been found to be positively associated with discretionary accruals (Menon and Williams 2004). We include our measures of significant changes in company financing (*FINANCE*) and whether the company entered into an acquisition during the period (*ACQ*) due to concerns over the effects of significant changes in capital structure on accounting accruals and because of the possible limitations of accruals management imposed by the structure of the balance sheet (Barton and Simko 2002). Prior research by Ashbaugh et al. (2003) and Chung and Kallapur (2003) provide evidence that significant changes in financing and entering into an acquisition are positively related to discretionary accruals. To control for prior performance we include lagged return on assets (*ROA*).

We initially run separate regressions for each of the three years using *DA* as the dependent variable, and then run another regression for the change in DA (ΔDA) from $t-1$ to $t+1$ as the dependent variable. The variable of interest is *HIRE*, which takes on a value of 1 for our sample CFO hiring firm, and 0 otherwise.[9] Panel A of Table 4 presents the regression results from the separate $t-1$, t, and $t+1$ time period analyses for the full sample of all firms with available control data in each period.[10] Consistent with the univariate analyses, after controlling for the other plausible factors associated with discretionary accruals, we find no significant differences between our HIRE firms and the other non-HIRE firms for DA in time t ($p = .912$), the initial year of the new CFO appointment. However, we find a significant ($p < .05$) positive relationship between DA and HIRE in time $t-1$, representing the former CFO's final full reporting year, and a significant ($p < .05$) negative DA relationship in time $t+1$, representing the new CFO's initial year of full reporting responsibility. In general, the control factors, except for the sign on *MVE*, have the expected signs and are significant at conventional levels.[11] In sum, these yearly multivariate results are consistent with the univariate results for levels of discretionary accruals and reflect a pattern of positive discretionary accruals in $t-1$ changing to negative accruals by $t+1$ for our sample of HIRE firms.

The regression for the change in discretionary accruals (ΔDA) over the $t-1$ to $t+1$ time period is presented in Panel B of Table 4. Because we are examining current discretionary accruals in our study, and reversals of these current accruals are more likely to take place within the ensuing period under study, we include the initial level of reported DA at $t-1$ in our multivariate analyses of ΔDA. Thus, when assessing changes in DA for our sample of 26,096 firms with all available data for the three-year time period, we include an additional variable, DA_{t-1}, into the model to control for the initial level of discretionary accruals exhibited by the firm at the start of the period.[12] We would expect there to be a negative relationship between initial levels of signed DA at $t-1$ and the change in DA over the ensuing two years. That is, firms reporting large positive DA at $t-1$ would be more

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[9] All non-indicator variable control variables are winsorized at the 1 and 99 percent levels.

[10] If we limit our yearly regressions to only the 26,096 firms that have complete data for all periods, then we obtain substantively the same results as those presented.

[11] We also perform tests for multicollinearity and do not find evidence of any significant effect on our results. Specifically, we calculated variance inflation factors for all coefficients for all regression models presented in the paper and find that the highest variance inflation factor is less than 2.0.

[12] Results of the models not including the initial level of discretionary accruals, or subsequently in our robustness tests, the initial level of performance-matched discretionary accruals, are substantively the same as those presented.

TABLE 4
Regression Results
Full Sample

Panel A: Discretionary Accruals (DA): By Year

Variables	**Exp. Sign**	**Time ($t-1$)** **(1)**	**Time (t)** **(2)**	**Time ($t+1$)** **(3)**
HIRE	+/?/−	0.0089**	0.0005	−0.0094**
		(0.034)	(0.912)	(0.040)
Controls				
MVE	−	0.0038***	0.0030***	0.0032***
		(0.000)	(0.000)	(0.000)
BM	−	−0.0021***	−0.0023***	−0.0022***
		(0.001)	(0.000)	(0.000)
DISTRESS	−	−0.0202***	−0.0186***	−0.0154***
		(0.000)	(0.000)	(0.000)
CFFO	−	−0.1911***	−0.1642***	−0.1459***
		(0.000)	(0.000)	(0.000)
GROWTH	−	−0.0038***	−0.0019***	−0.0017***
		(0.000)	(0.003)	(0.009)
FINANCE	+	0.0158***	0.0149***	0.0129***
		(0.000)	(0.000)	(0.000)
ACQ	+	0.0165***	0.0158***	0.0152***
		(0.000)	(0.000)	(0.000)
ROA_{t-1}	+	0.0029***	0.0019***	0.0013***
		(0.000)	(0.000)	(0.000)
CONSTANT	?	−0.0830***	−0.0742***	−0.0654***
		(0.000)	(0.000)	(0.000)
n =		32,037	33,157	33,640
Prob > F		0.000***	0.000***	0.000***
R^2		0.169	0.155	0.147

Panel B: Changes in Discretionary Accruals (DA): Time ($t+1$) minus Time ($t-1$)

Variables	**Exp. Sign**	**ΔDA** **(1)**
HIRE	−	−0.0128**
		(0.015)
Controls		
DA_{t-1}	−	−0.1444***
		(0.000)
Δ*MVE*	−	0.0165***
		(0.000)
Δ*BM*	−	−0.0027***
		(0.000)
Δ*DISTRESS*	−	−0.1997***
		(0.000)
Δ*CFFO*	−	−0.2511***
		(0.000)

(*continued on next page*)

TABLE 4 (Continued)

Panel B: Changes in Discretionary Accruals (DA): Time ($t+1$) minus Time ($t-1$) (Continued)

Variables	Exp. Sign	ΔDA (1)
Δ*GROWTH*	−	0.0046*** (0.000)
Δ*FINANCE*	+	0.0058*** (0.000)
Δ*ACQ*	+	−0.0028* (0.099)
ΔROA_{t-1}	+	0.0311*** (0.000)
CONSTANT	?	−0.0024*** (0.006)
n =		26,096
Prob > F		0.000***
R2		0.294

*, **, *** Denotes significance at the 10%, 5%, and 1% levels, respectively.

Multivariate regressions models for the full sample of HIRE and Non-HIRE companies in the period 1994–2000. The HIRE sample includes 712 companies that appointed a new CFO. Time (t) is the year of the new CFO appointment. Dependent variables are discretionary accruals (DA) and in changes in discretionary accruals (ΔDA).

p-values are in parentheses. p-values for independent variable t-tests were calculated using two-tailed tests.

Control variables in Panel A are defined as: *MVE* = log of the market value of equity; *BM* = book-to-market equity ratio; *DISTRESS* = financial distress measure (calculated from Zmijewski 1984); *CFFO* = cash flow from operations divided by total assets; *GROWTH* = sales growth rate; *FINANCE* = 1 if number of o/s shares increased by at least 10 percent or long-term debt increased by at least 20 percent during the year; *ACQ* = 1 if the company engaged in an acquisition, and HIRE = 1 if company hired a CFO, 0 otherwise; ROA_{t-1} = return on assets from the prior year-end.

Control variables in Panel B are defined as: DA_{t-1} = DA from $time_{t-1}$, and all others are changes in the control variables as defined in Panel A.

likely to report larger reductions in levels of signed DA in the ensuing years than firms reporting lower or negative DA in $t-1$.

As reported in Panel B of Table 4, our ΔDA model results indicate, after controlling for initial levels of accruals and for changes in the other control factors, a significant negative coefficient on HIRE. This result confirms our year-by-year analyses and provides further support that our sample of CFO-hiring firms exhibited significant reductions in discretionary accruals compared to other non-HIRE firms over the period, even after controlling for changes in factors associated with discretionary accruals. These multivariate results suggest that, after controlling for other factors, our HIRE sample firms reduced DA an average of an additional $12.22 million over the three-year reporting period surrounding a CFO change compared to our non-HIRE control firms. Accordingly, our univariate and multivariate results provide consistent support for H_1 regarding the reduction in accruals surrounding the appointment of a new individual to the CFO position.

Effects of Concurrent CEO Turnover

Prior research has documented significant changes in financial reporting surrounding a CEO appointment. In order to ensure that our results are not confounded by a contemporaneous CEO change during our CFO hire examination period, we identify cases in which our HIRE firm concurrently appoints a new CEO during the $t-1$ to $t+1$ examination period.

For our sample of 712 HIRE firms we find that 322 (45.2 percent) firms also appoint a new CEO during the $t-1$ to $t+1$ period surrounding the hire of a new CFO.[13] Accordingly, we assess whether changes in abnormal accruals are associated with contemporaneous CEO appointments in our examination period surrounding CFO appointments by separating our 712 HIRE sample firms into the 390 firms without a concurrent CEO appointment and the 322 firms with a contemporaneous CEO hire. Results of the separate subsample univariate analyses are presented in the second sections of Panels A and B of Table 3. The separate subsample multivariate analyses are presented in Table 5.

As noted in Table 3, our 390 HIRE firms without a contemporaneous CEO appointment exhibit a similar pattern of year-end DA levels and changes in DA when compared to the non-HIRE firms as did our comparisons of the entire HIRE sample. However, our 322 HIRE firms that had a CEO hiring in our CFO turnover window present a similar yearly and directional pattern of reduction in DA as the aggregate HIRE sample, although differences for this subsample are generally not significantly different from the non-HIRE control sample. Specifically, the subsample without a CEO hire exhibited ΔDA of −3.60 percent, while the subsample with a concurrent CEO appointment had a mean ΔDA of only −2.22 percent. Compared to the ΔDA of −0.86 percent for the control sample of non-HIRE firms, the ΔDA for the subsample of HIRE firms without CEO turnover is significant different ($p < .01$), while the ΔDA for the subsample of HIRE firms with a concurrent appointment of a new CEO is not ($p > .10$).

These subsample univariate results are also mirrored in the separate multivariate regressions presented in Table 5. Panel A presents the results of the multivariate regressions presented earlier and including only the 390 HIRE firms without a concurrent CEO appointment and the control samples of non-HIRE control firms. Results of the yearly regressions for DA and for ΔDA are very similar to those presented in Table 4 when using the entire HIRE sample. In this subsample, the HIRE coefficient is positive and significant ($p < .05$), not significant, and negative and significant ($p < .10$) for the $t-1$, t, and $t+1$ periods, respectively. Similarly, the ΔDA regression produces a HIRE coefficient that is negative and significant ($p < .01$) indicating that the 390 HIRE firms without CEO turnover reported higher reductions in DA than did the control sample of non-HIRE firms. In contrast, however, none of the regressions in Panel B for the HIRE subsample with concurrent CEO turnover and the non-HIRE control samples indicate a significant difference in yearly DA levels or in ΔDA compared to the non-HIRE control firms.

To further examine the possible effect of a concurrent CEO hiring on our results, we test to see whether the timing of a new CEO appointment around the time a new CFO is hired makes a difference with respect to the changes in DA reported by our sample HIRE firms. For example, CEOs appointed prior to a new CFO would most likely have some role in the CFO selection. CFOs selected by the CEO might exhibit different changes in accruals compared to CFOs appointed immediately prior to the new CEO. Accordingly, we add indicator variables to our multivariate models in Table 4 for whether the CEO was appointed before the CFO (CEO_{pre}) or after the CFO (CEO_{post}). Results of these expanded models produce non-significant coefficients ($p > .10$) on both CEO indicator variables and the HIRE coefficients remain essentially the same as presented.[14] Similarly, if we separately

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[13] Fourteen firms in our sample had two concurrent CEO and CFO appointments during our 1994–2000 examination period. Deleting these 14 firms does not change the results of our analyses.

[14] Using three separate indicator variables for CEO appointments in each of the three years in our examination period produces similar results.

TABLE 5
Regression Results
Without CEO Turnover and With CEO Turnover

Panel A: HIRE Subsample without CEO Turnover (n = 390)

Variables	Exp. Sign	Time ($t-1$) (1)	Time (t) (2)	Time ($t+1$) (3)	Variables	Exp. Sign	ΔDA (1)
HIRE	+/?/−	0.0118**	0.0050	−0.0109*	**HIRE**	−	−0.0182***
w/CEO = 0		(0.038)	(0.392)	(0.076)	w/CEO = 0		(0.010)
Controls							
					DA_{t-1}	−	−0.1436***
							(0.000)
MVE	−	0.0038***	0.0030***	0.0032***	*ΔMVE*	−	0.0165***
		(0.000)	(0.000)	(0.000)			(0.000)
BM	−	−0.00133*	−0.0023***	−0.0021***	*ΔBM*	−	−0.0026***
		(0.082)	(0.000)	(0.000)			(0.000)
DISTRESS	−	−0.0202***	−0.0186***	−0.0154***	*ΔDISTRESS*	−	−0.2006***
		(0.000)	(0.000)	(0.000)			(0.000)
CFFO	−	−0.1913***	−0.1647***	−0.1461***	*ΔCFFO*	−	−0.2514***
		(0.000)	(0.000)	(0.000)			(0.000)
GROWTH	−	−0.0038***	−0.0019***	−0.0017***	*ΔGROWTH*	−	0.0047***
		(0.000)	(0.002)	(0.009)			(0.000)
FINANCE	+	0.0157***	0.0148***	0.0129***	*ΔFINANCE*	+	0.0057***
		(0.000)	(0.000)	(0.000)			(0.000)
ACQ	+	0.0162***	0.0158***	0.0151***	*ΔACQ*	+	−0.0032*
		(0.000)	(0.000)	(0.000)			(0.069)
ROA_{t-1}	+	0.0029***	0.0019***	0.0012***	ΔROA_{t-1}	+	0.0310***
		(0.000)	(0.000)	(0.000)			(0.000)
CONSTANT	?	−0.0830***	−0.0742***	−0.0653***	*CONSTANT*	?	−0.0023***
		(0.000)	(0.000)	(0.000)			(0.007)
n =		31,715	32,835	33,318	n =		25,774
Prob > F		0.000***	0.000***	0.000***	Prob > F		0.000***
R^2		0.169	0.155	0.147	R^2		0.294

Panel B: HIRE Subsample with CEO Turnover (n = 322)

Variables	Exp. Sign	Time ($t-1$) (1)	Time (t) (2)	Time ($t+1$) (3)	Variables	Exp. Sign	ΔDA (1)
HIRE	+/?/−	0.0055	−0.0050	−0.0074	**HIRE**	−	−0.0063
w/CEO = 1		(0.375)	(0.436)	(0.268)	w/CEO = 1		(0.422)
Controls							
					DA_{t-1}	−	−0.1438***
							(0.000)
MVE	−	0.0038***	0.0031***	0.0032***	*ΔMVE*	−	0.0165***
		(0.000)	(0.000)	(0.000)			(0.000)
BM	−	−0.0012	−0.0024***	−0.0023***	*ΔBM*	−	−0.0027***
		(0.111)	(0.000)	(0.000)			(0.000)
DISTRESS	−	−0.0201***	−0.0186***	−0.0154***	*ΔDISTRESS*	−	−0.2002***
		(0.000)	(0.000)	(0.000)			(0.000)

(*continued on next page*)

TABLE 5 (Continued)

Panel B: HIRE Subsample with CEO Turnover (n = 322) (Continued)

Variables	Exp. Sign	Time ($t-1$) (1)	Time (t) (2)	Time ($t+1$) (3)	Variables	Exp. Sign	ΔDA (1)
CFFO	–	−0.1911***	−0.1642***	−0.1459***	Δ*CFFO*	–	−0.2514***
		(0.000)	(0.000)	(0.000)			(0.000)
GROWTH	–	−0.0038***	−0.0019***	−0.0017***	Δ*GROWTH*	–	0.0046***
		(0.000)	(0.002)	(0.009)			(0.000)
FINANCE	+	0.0155***	0.0148***	0.0126***	Δ*FINANCE*	+	0.0058***
		(0.000)	(0.000)	(0.0000			(0.000)
ACQ	+	0.0163***	0.0161***	0.0153***	Δ*ACQ*	+	−0.0030*
		(0.000)	(0.000)	(0.000)			(0.084)
ROA_{t-1}	+	0.0029***	0.0019***	0.00127***	ΔROA_{t-1}	+	0.0310***
		(0.000)	(0.000)	(0.000)			(0.000)
CONSTANT	?	0.0830***	−0.0744***	−0.0653***	*CONSTANT*	?	−0.0024***
		(0.000)	(0.000)	(0.000)			(0.007)
n =		31,647	32,767	33,250	n =		25,706
Prob > F		0.000***	0.000***	0.000***	Prob > F		0.000***
R^2		0.169	0.155	0.147	R^2		0.293

*, **, *** Denotes significance at the 10%, 5%, and 1% levels, respectively.

Multivariate regression models for the subsamples of HIRE companies with and without concurrent CEO turnover along with all Non-HIRE companies in the period 1994–2000. The HIRE sample includes companies that appointed a new CFO. Time (t) is the year of the new CFO appointment. Dependent variables are discretionary accruals (DA) and in changes in discretionary accruals (ΔDA).

p-values are in parentheses. p-values for independent variable t-tests were calculated using two-tailed tests.

Control variables in the DA regressions are defined as: *MVE* = log of the market value of equity; *BM* = book-to-market equity ratio; *DISTRESS* = financial distress measure (calculated from Zmijewski 1984); *CFFO* = cash flow from operations divided by total assets; *GROWTH* = sales growth rate; *FINANCE* = 1 if number of o/s shares increased by at least 10 percent or long-term debt increased by at least 20 percent during the year; *ACQ* = 1 if the company engaged in an acquisition, and HIRE = 1 if company hired a CFO, 0 otherwise; ROA_{t-1} = return on assets from the prior year-end.

Control variables in the ΔDA regressions are defined as: DA_{t-1} = DA from time$_{t-1}$, and all others are changes in the control variables as defined in the DA regressions.

assess the 188 CEO_{pre} firms and the 144 CEO_{post} firms against our control sample of non-HIRE firms in our multivariate regressions (as in Table 4) we obtain nonsignificant results on the CFO HIRE subsample indicator variables in all analyses.

These additional analyses provide consistent evidence that the appointment of a new CEO around the time a new CFO is hired does not appear to be confounding our results. In fact, we find that cases were there are no concurrent CEO appointments in our three-year examination window surrounding a new CFO result in the most significant reductions in DA compared to our control sample and provide the strongest evidence of the influence of the appointment of a new CFO on changes in the company's reported financial results. Our weak results of joint CEO and CFO turnover, however, may be due to the fact that all types of CEO changes (e.g., retirements, firings, internal appointments, external hires) are included in the analyses.

Univariate Results—H2

Our second hypothesis addresses whether a new CFO coming from outside the company results in a higher reductions in discretionary accruals than when an individual is appointed from within the firm. With respect to changes in reported discretionary accounting accruals, we would expect that promoting or appointing an individual from within the company would result in relatively smaller changes in financial reporting than hiring a new CFO from outside the company. Accordingly, we parse our sample of 712 HIRE firms based on the new CFO's former employer. In our sample of 712 HIRE firms we find that 574 (80.6 percent) were appointed from outside the company (EXTERNAL), while 138 (19.4 percent) were appointed to CFO from other positions from within the firm (INTERNAL). The last section of Panels A and B of Table 3 present the univariate results of comparing these subsamples of our HIRE firms to our sample of non-HIRE control firms.

As shown in Table 3, the results regarding the INTERNAL CFO-hiring firms are very different from that exhibited by the EXTERNAL hiring firms. Specifically, our subsample of EXTERNAL CFO-hiring firms report significantly higher DA in $t-1$ ($p < .01$), no different DA in t, and significantly lower DA in $t+1$ ($p < .01$) compared to the non-HIRE control firms. These results are similar to our aggregate HIRE sample results presented earlier. However, the INTERNAL CFO-hiring firms report very similar DA at the end of each of the three years as does our control non-HIRE group. While we might expect the large group of EXTERNAL CFO-hiring firms to report similar DA levels as our aggregate HIRE sample, our results for the INTERNAL CFO-hiring firms is substantially different from our aggregate HIRE sample results and are extraordinarily similar to the non-HIRE control firms. These overall patterns of DA and specifically the changes in DA (ΔDA) surrounding the CFO appointment are explicitly tested and presented in Panel B of Table 3.

Results of the univariate tests for ΔDA from $t-1$ to $t+1$ indicate that the change in DA for the EXTERNAL hiring firms of -3.47 percent is significantly different from the change of -0.86 percent for the non-HIRE control firms ($p < .01$). However, the change in DA for the INTERNAL hiring firms of -0.94 percent is not significantly different from that of the non-HIRE control firms ($p > .10$). These univariate results indicate that firms hiring an EXTERNAL CFO report significantly larger reductions in DA surrounding this new appointment compared to non-hiring firms, while firms appointing INTERNAL CFOs do not exhibit any significant differences in levels or ΔDA compared to our control sample of non-HIRE firms.

Multivariate Results—H2

In order to assess whether our univariate results with respect to changes in DA for EXTERNAL and INTERNAL CFO hires are robust to other factors associated with levels of DA, we re-estimate our multivariate models from Table 4 using these two separate subsamples. Results of the EXTERNAL hire subsample are presented in Panel A of Table 6 and the results of the INTERNAL hire subsample are presented in Panel B.

Consistent with the univariate results, and with our second hypothesis, the results of comparing the EXTERNAL firms with the non-HIRE control firms reveals that the yearly regressions for DA and for the ΔDA regression are very similar to those presented in Table 4 when using the entire HIRE sample. Using this subsample of EXTERNAL firms, Panel A indicates that the HIRE coefficient is positive and significant ($p < .05$), not significant, and negative and significant ($p < .05$) for the $t-1$, t, and $t+1$ periods, respectively. Similarly, the ΔDA regression produces a HIRE coefficient that is also negative and significant ($p < .01$) indicating that the EXTERNAL CFO-hiring firms reported larger reductions in DA over the years surrounding the CFO appointment than did the control sample of non-HIRE firms. These EXTERNAL subsample regression results are very consistent with the overall regression results for the entire HIRE sample. However, as anticipated from the univariate analyses, none of the regressions in Panel B for the HIRE subsample appointing

TABLE 6
Regression Results
External Hires and Internal Hires

Panel A: External HIRE (n = 574)

Variables	Exp. Sign	Time ($t-1$) (1)	Time (t) (2)	Time ($t+1$) (3)	Variables	Exp. Sign	ΔDA (1)
HIRE	+/?/−	0.0092**	−0.0001	−0.0126**	**HIRE**	−	−0.0162***
		(0.048)	(0.979)	(0.013)			(0.006)
Controls							
					DA_{t-1}	−	−0.1442***
							(0.000)
MVE	−	0.0038***	0.0031***	0.0032***	ΔMVE	−	0.0165***
		(0.000)	(0.000)	(0.000)			(0.000)
BM	−	−0.0012*	−0.0023***	−0.0022***	ΔBM	−	−0.0026***
		(0.095)	(0.000)	(0.000)			(0.000)
DISTRESS	−	−0.0202***	−0.0186***	−0.0154***	$\Delta DISTRESS$	−	−0.2008***
		(0.000)	(0.000)	(0.000)			(0.000)
CFFO	−	−0.1912***	−0.1644***	−0.1458***	$\Delta CFFO$	−	−0.2511***
		(0.000)	(0.000)	(0.000)			(0.000)
GROWTH	−	−0.0038***	−0.0019***	−0.0017***	$\Delta GROWTH$	−	0.0045***
		(0.000)	(0.0030	(0.009)			(0.000)
FINANCE	+	0.0158***	0.0148***	0.0128***	$\Delta FINANCE$	+	0.0058***
		(0.000)	(0.000)	(0.000)			(0.000)
ACQ	+	0.0164***	0.0159***	0.0152***	ΔACQ	+	−0.0029*
		(0.000)	(0.000)	(0.000)			(0.097)
ROA_{t-1}	+	0.0029***	0.0019***	0.0012***	ΔROA_{t-1}	+	0.0309***
		(0.000)	(0.000)	(0.000)			(0.000)
CONSTANT	?	−0.0831***	−0.0742***	−0.0654***	*CONSTANT*	?	−0.0024***
		(0.000)	(0.000)	(0.000)			(0.006)
n =		31,899	33,019	33,502	n =		25,958
Prob > F		0.000***	0.000***	0.000***	Prob > F		0.000***
R^2		0.169	0.155	0.147	R^2		0.294

Panel B: Internal HIRE (n = 138)

Variables	Exp. Sign	Time ($t-1$) (1)	Time (t) (2)	Time ($t+1$) (3)	Variables	Exp. Sign	ΔDA (1)
HIRE	+/?/−	0.0076	0.0029	0.0041	**HIRE**	−	0.0012
		(0.424)	(0.762)	(0.690)			(0.918)
Controls							
					DA_{t-1}	−	−0.1433***
							(0.000)
MVE	−	0.0038***	0.0031***	0.0031***	ΔMVE	−	0.0165***
		(0.000)	(0.000)	(0.000)			(0.000)
BM	−	−0.0012*	−0.0024***	−0.0022***	ΔBM	−	−0.0026***
		(0.099)	(0.000)	(0.000)			(0.000)
DISTRESS	−	−0.0201***	−0.0186***	−0.0154***	$\Delta DISTRESS$	−	−0.2000***
		(0.000)	(0.000)	(0.000)			(0.000)

(*continued on next page*)

TABLE 6 (Continued)

Panel B: Internal HIRE (n = 138) (Continued)

Variables	Exp. Sign	Time $(t-1)$ (1)	Time (t) (2)	Time $(t+1)$ (3)	Variables	Exp. Sign	ΔDA (1)
CFFO	–	−0.1912***	−0.1645***	−0.1462***	Δ*CFFO*	–	−0.2517***
		(0.000)	(0.000)	(0.000)			(0.000)
GROWTH	–	−0.0037***	−0.0020***	−0.0017***	Δ*GROWTH*	–	0.0048***
		(0.000)	(0.000)	(0.009)			(0.000)
FINANCE	+	0.0154***	0.0148***	0.0128***	Δ*FINANCE*	+	0.0057***
		(0.000)	(0.000)	(0.000)			(0.000)
ACQ	+	0.0160***	0.0159***	0.0152***	Δ*ACQ*	–	−0.0033*
		(0.000)	(0.000)	(0.000)			(0.057)
ROA_{t-1}	+	0.0029***	0.0019***	0.0012***	ΔROA_{t-1}	–	0.0310***
		(0.000)	(0.000)	(0.000)			(0.000)
CONSTANT	?	−0.0830***	−0.0743***	−0.0652***	*CONSTANT*	?	−0.0024***
		(0.000)	(0.000)	(0.000)			(0.006)
n =		31,463	32,583	33,066	n =		25,522
Prob > F		0.000***	0.000***	0.000***	Prob > F		0.000***
R^2		0.168	0.155	0.147	R^2		0.293

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*, **, *** Denotes significance at the 10%, 5%, and 1% levels, respectively.

Multivariate regressions models for the subsamples of internal and external HIRE companies and all Non-HIRE companies in the period 1994–2000. The HIRE subsamples include companies that appointed a new CFO. Time (t) is the year of the new CFO appointment. Dependent variables are discretionary accruals (DA) and in changes in discretionary accruals (ΔDA).

p-values are in parentheses. p-values for independent variable t-tests were calculated using two-tailed tests. Control variables in the DA regressions are defined as: *MVE* = log of the market value of equity; *BM* = book-to-market equity ratio; *DISTRESS* = financial distress measure (calculated from Zmijewski 1984); *CFFO* = cash flow from operations divided by total assets; *GROWTH* = sales growth rate; FINANCE = 1 if number of o/s shares increased by at least 10 percent or long-term debt increased by at least 20 percent during the year; *ACQ* = 1 if the company engaged in an acquisition, and HIRE = 1 if company hired a CFO, 0 otherwise; ROA_{t-1} = return on assets from the prior year-end. Control variables in the ΔDA regressions are defined as: DA_{t-1} = DA from time_{t-1}, and all others are changes in the control variables as defined in the DA regressions.

an INTERNAL CFO result in a significant difference for any of the yearly DA regressions or in the ΔDA regression, indicating the yearly DA levels and ΔDA for firms making internal appointments to CFO are not significantly different from the non-HIRE control firms.

In sum, our univariate and multivariate results consistently indicate that the firms hiring an EXTERNAL CFO significantly reduced signed DA surrounding the new CFO hire, while firms appointing an INTERNAL person to CFO show no such reductions to DA surrounding that appointment. Accordingly, we find strong support for our second hypothesis that external CFO appointments would result in significantly larger reductions in reported DA than would be exhibited by firms appointing a new CFO from within their corporate ranks.

We then examine our EXTERNAL hire subsample for a "revolving door" effect on changes in discretionary accruals by identifying whether the new CFO was hired directly from the company's external CPA firm. Recent studies (Dowdell and Krishnan 2004; Geiger et al. 2005) have examined accounting accruals surrounding the hire of an individual directly from the company's external audit firm and have reached conflicting conclusions. Accordingly, in order to assess any possible "revolving door" hiring effect within our sample, we expand our multivariate ΔDA model by adding an indicator variable for when the new CFO comes either directly from the firm's current external auditor (*CPAFIRM*; n = 18)—the "revolving door" hires—or from any CPA firm (*CPAEXP*; n = 42). Including the *CPAEXP* indicator variable allows us to examine our sample for any general CPA experience effects. Results of these expanded regressions separately including one indicator variable, and including both indicators (not tabulated here) indicate nonsignificant coefficient estimates for both of these two indicator variables ($p > .25$) in all expanded regressions, yet the HIRE indicator coefficients remain negative and significant ($p < .01$) in all analyses. Thus, our results do not indicate any difference in changes in DA based on whether the individual came directly from a CPA firm, or directly from the company's current external audit firm. Accordingly, we find no "revolving door" effect for companies in our sample hiring CFOs directly from audit firms, or directly from their current external audit firm. These results are consistent with those of Geiger et al. (2005) in that companies hiring new CFOs directly from their CPA firms exhibit similar changes in discretionary accruals as other non-revolving door CFO hires.

IV. FURTHER TESTS

Previous researchers (Dechow et al. 1995; Ashbaugh et al. 2003; Kothari et al 2005) have noted that accruals estimation models such as the cross-sectional Jones (1991) model used in our study may be misspecified when applied to companies that are significantly different with respect to financial performance. Kothari et al. (2005) argues that the misspecification problem is mitigated by performance-matching when estimating levels of abnormal accruals. Accordingly, while we include *ROA* in our multivariate models to provide control for any financial performance effects, in order to mitigate any possible residual effect of company financial performance on our discretionary accruals estimations, we construct a yearly performance-matched subsample of 712 of the non-HIRE companies as an additional control sample. To identify our performance-matched sample we use the matching procedure outlined in Kothari et al. (2005) where each HIRE company is matched with a non-HIRE company from the same two-digit SIC code that has the closest *ROA* for the year.[15] Since we are examining multiple years, we perform this matching procedure separately for each of the years $t-1$, t, and $t+1$surrounding the appointment of the new CFO.[16] Thus, our measure of performance-adjusted discretionary accruals (*PMDA*) becomes $PMDA = DA_h - DA_m$, where DA_h represents the discretionary accruals of the hth HIRE firm, and DA_m represents the discretionary accruals of the corresponding performance-matched non-HIRE firm.

Using *PMDA* in our tests instead of DA, and $\Delta PMDA$ (i.e., changes in *PMDA*) in place of ΔDA, does not substantively change our results. For example, in examining the signed *PMDA* for time $t-1$, we find that the mean of 1.37 percent for *PMDA* is significantly different from zero ($p < .01$), indicating that the HIRE firms used significantly more income-increasing accruals than their performance-matched non-HIRE firms in $t-1$. However, as in the *DA* analyses, this significant difference in level of *PMDA* disappears in time t (mean of only 0.40 percent), and then becomes negative and significant in time $t+1$ (mean of -1.05 percent, significant at $p < .01$). Further, these reductions in *PMDA* for our sample of HIRE firms over the three years is also significantly greater than for all 25,399 non-HIRE firms with available data for the same period ($p < .01$). Additionally, the results of our multivariate tests using *PMDA* and Δ*PMDA* (after removing *ROA* from the models) produces similar results as those presented in the paper. Thus, our results do not appear to be overly sensitive to the performance characteristics of our HIRE firms and are generally robust across this different specification of discretionary accruals.

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[15] Potential matching companies were then further screened to ensure that they did not hire a new CFO in our test period (i.e., from $t-1$ to $t+1$).

[16] Alternatively, conducting the performance-matching of HIRE firms to non-HIRE firms only at time $t-1$ and tracking changes in these performance-adjusted accruals across the three-year period yields similar results.

In order to ensure that our results are robust across different segments of our HIRE firms, we perform several additional analyses. First, we assess whether our findings are not affected by firm size. While we scale for firm size in our calculation of discretionary accruals and control for firm size in our regressions, smaller firms may be differentially affected by CFO turnover. Accordingly, if we add the log of the market value of equity to our ΔDA regression model ($LNMVE_t$) measured at time t, similar to Menon and Williams (2004), we find a negative and significant ($p < .05$) coefficient on this size measure and substantively the same results on the HIRE variable (negative and significant at $p < .05$) and the other control variables. Thus, while larger firms appear to report significantly greater reductions in DA surrounding a CFO change, it does not appear that the reduction in DA for our sample of HIRE firms is significantly effected by firm size.

We then assess whether our results are stable over our multiyear time period or whether they are driven by more recent CFO hiring events. Accordingly, we add a dummy variable for time based on year of hire ($\text{YEAR}_{1998-2000}$; 1 if year > 1998) in order to assess whether the change in DA for our sample of reporting firms has substantially changed over time by comparing the earlier years to the later years in our 1994–2000 examination period. Inclusion of the $\text{YEAR}_{1998-2000}$ indicator variable to the ΔDA regression results in a negative and significant coefficient estimate ($p < .05$). Thus, reductions in DA surrounding the hire of a new CFO appear to be more salient in the later years of our sample period. However, inclusion of the indicator variable does not change the results with respect to the HIRE coefficient (negative and significant at $p < .01$) and the other control variables, indicating that while there may be differences across time subperiods, these differences do not significantly effect our aggregate results regarding reduced DA for our HIRE firms surrounding the CFO appointment.

Last, we assess whether our significant results for changes in signed DA is consistent throughout our examination window surrounding the CFO turnover event. More specifically, we examine whether the DA reductions occur ratably over the three-year examination window, or whether our aggregate ΔDA reductions are largely driven by the high level of discretionary accruals for our HIRE firms in $t-1$ that result in relatively larger reductions in $t-1$ to t. Table 7 presents the regression results if we separately examine changes in accruals for the period immediately prior to appointing a new CFO to the first year of reported results ($t-1$ to t), and the period from appointment to the immediately following period (t to $t+1$). As in the earlier regressions, we control for initial levels of DA for both of these analyses (i.e., levels at $t-1$ or t, respectively). Our regression results for the full HIRE sample in Panel A indicate that changes in DA are negative but not significant for the $t-1$ to t period, and also for the t to $t+1$ period ($p > .10$). These aggregate results suggest that the changes in DA for our HIRE firms are relatively consistent over our examination period surrounding the appointment of a new CFO and that our ΔDA results are not driven by DA changes in any one year. In fact, we find no significant reduction in DA for either of the separate two year analyses. These results suggest that it is important to examine the entire turnover time period surrounding the CFO change and not just the reported DA results for any one year, or the change in reported DA over a single year that may prove insightful.

In order to determine if these yearly ΔDA results are consistent across the subsamples of HIRE firms examined in the study, we also examined the two subsamples in our earlier

TABLE 7
Additional Analyses: One Year Changes in Discretionary Accruals

	Panel A		Panel B		Panel C	
	All HIRES (n = 712)		HIRES without CEO turnover (n = 390)		External HIRES (n = 574)	
Time	(t) − (t−1)	(t+1) − (t)	(t) − (t−1)	(t+1) − (t)	(t) − (t−1)	(t+1) − (t)
Variables	ΔDA (1)	ΔDA (2)	ΔDA (1)	ΔDA (2)	ΔDA (1)	ΔDA (2)
HIRE	−0.0078	−0.0054	−0.0050	−0.0111*	−0.0085	−0.0080
	(0.113)	(0.263)	(0.446)	(0.091)	(0.117)	(0.143)
Controls						
$DA_{t-1,t}$	−0.1285***	−0.1814***	−0.1279***	−0.1803***	−0.1282***	−0.1811***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
ΔMVE	0.0152***	0.0161***	0.0154***	0.0161***	0.0153***	0.0161***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
ΔBM	−0.0015	0.0001	−0.0014	0.0001	−0.0015	0.0001
	(0.100)	(0.885)	(0.135)	(0.809)	(0.106)	(0.910)
$\Delta DISTRESS$	−0.2197***	−0.1924***	−0.2190***	−0.1928***	−0.2201***	−0.1930***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$\Delta CFFO$	−0.3311***	−0.2927***	−0.3314***	−0.2938***	−0.3309***	−0.2925***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$\Delta GROWTH$	0.0022***	0.0024***	0.0022**	0.0024***	0.0022**	0.0024***
	(0.010)	(0.006)	(0.012)	(0.006)	(0.011)	(0.006)
$\Delta FINANCE$	0.0044***	0.0045***	0.0044***	0.0044***	0.0043***	0.0045***
	(0.001)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
ΔACQ	−0.0022	−0.0045***	−0.0024	−0.0044**	−0.0022	−0.0044**
	(0.195)	(0.010)	(0.169)	(0.011)	(0.213)	(0.011)
ΔROA_{t-1}	0.0043**	0.0001***	0.0045**	0.0001***	0.0043**	0.0001***
	(0.038)	(0.000)	(0.032)	(0.000)	(0.041)	(0.000)
CONSTANT	0.0008	−0.0008	0.0008	−0.0008	0.0008	−0.0008
	(0.320)	(0.302)	(0.330)	(0.305)	(0.325)	(0.308)
n =	26,096	26,096	25,774	25,774	25,958	25,958
Prob > F	0.000***	0.000***	0.000***	0.000***	0.000***	0.000***
R^2	0.330	0.320	0.330	0.320	0.330	0.320

*, **, *** Denotes significance at the 10%, 5%, and 1% levels, respectively.

Multivariate regression models for the sample of all HIRE companies or subsamples of HIRE companies and all non-HIRE companies in the period 1994–2000. Dependent variables are changes in discretionary accruals (ΔDA) over a one-year period. Time (*t*) is the year of the new CFO appointment.

p-values are in parentheses. p-values for independent variable t-tests were calculated using two-tailed tests. Control variables are defined in Table 4.

analyses where we find similar significant reductions in DA across the $t-1$ to $t+1$ period. Results of the examination of the non-CEO turnover subsample analyses are presented in Panel B of Table 7, and results of the examination of the EXTERNAL CFO-hiring firms subsample are presented in Panel C. As noted in both panels, the turnover subperiod analyses for these two HIRE subsamples also suggest that the reductions in DA occur relatively evenly over the CFO turnover window and that our ΔDA results are not driven by changes in only one of the two years. Thus, we obtain similar results for these subsamples as found for the aggregate HIRE sample. Not only do our results provide evidence that the change in DA surrounding the hire of a new CFO is not simply the result of accruals reductions in any one year, but we also find strong evidence that the decreases in accruals in the $t-1$ to t period and the t to $t+1$ period, while not significant by themselves, in the aggregate indicate that firms appointing a new CFO report significant reductions in DA from the last year under the former CFO to the first full reporting year under the new CFO.

V. CONCLUSIONS

In this study we assess the relationship between appointing a new CFO and corporate financial reporting by examining the changes in a firm's reported discretionary accounting accruals surrounding the new appointment. Specifically, we assess yearly levels and changes in levels of discretionary accruals from the year preceding the appointment of a new CFO to the year after their appointment, representing the time period under the full purview of the former CFO to the first year the new CFO had full responsibility for all company financial information.

Using the cross-sectional Jones (1991) accruals model, as well as the performance-matching procedure advocated in Kothari et al. (2005) as reported in our "Further Tests" section, and using multivariate regressions to provide additional control for other accrual-related factors, we find that, compared to non-hiring firms, discretionary accruals are significantly reduced surrounding the arrival of a new CFO for our sample of 712 CFO-hiring firms. Further, we find that our results are not confounded by the appointment of a new CEO during our CFO turnover examination period. In fact we find that firms not appointing a new CEO during our examination period report significant reductions in discretionary accruals while firms concurrently appointing a new CEO exhibit no significant reduction in discretionary accruals compared to non-hiring firms. However, as noted earlier, our results regarding joint CEO and CFO turnover may be due to the fact that we include all types of CEO changes (e.g., retirements, firings, internal appointments, external hires) in our analyses.

Our results also indicate that companies appointing individuals to the position of CFO that come from external sources report reductions in discretionary accruals, while companies appointing CFOs from within do not report similar reductions. We also find that firms hiring CFOs directly from their external audit firm, or from any other CPA firm, do not report any differences in the pattern of discretionary accruals reductions than other CFO appointments. These nonsignificant "revolving door" results are consistent with the findings of Geiger et al. (2005) who found no differences in the changes in absolute and nonoperating accruals for revolving-door hires compared to their control samples of non-revolving-door hires.

Our study contributes to the corporate governance and discretionary accruals literatures by documenting the consistent reductions in discretionary accruals reported by firms surrounding this important corporate hiring event. While prior research has documented the effects of CEO hiring on a company's reported financial results, our study extends this literature and presents evidence that changes in CFOs are associated with significant reductions in a company's reported discretionary accruals. Our results are also congruent with the recent perceptions of regulators in the U.S., as reflected in the new CFO and CEO reporting requirements in the Sarbanes-Oxley Act, that individual CFOs are very influential players in the corporate reporting arena. While they are not generally involved in profit-directed activities, our findings suggest that a change in a firm's CFO significantly affects the firm's reported financial results.

While we examine changes in discretionary accounting accruals, investigating other measures of reported financial performance surrounding this significant corporate event would also be fruitful areas for future research. Examining other measures of firm performance surrounding an appointment of a new CFO would also provide insight into the changes brought about under the new CFO. Further, an interesting question raised by our study is whether the outgoing CFOs time their departures to coincide with reporting favorable financial results through the aggressive use of discretionary accounting accruals. Assessing this possibility would lend additional evidence regarding the possible reasons for the changes in discretionary accruals surrounding the appointment of a new CFO, as found in this study.

REFERENCES

Aboody, D. and R. Kasznik. 2000. CEO stock option awards and the timing of corporate voluntary disclosures. *Journal of Accounting and Economics* 29: 73–100.

Ashbaugh, H., R. LaFond, and B. W. Mayhew. 2003. Do nonaudit services compromise auditor independence? Further evidence. *The Accounting Review* 78 (July): 611–640.

Barton, J., and P. J. Simko. 2002. The balance sheet as an earnings management constraint. *The Accounting Review* 77 (Supplement): 1–27.

Bartov, E., F. A. Gul, and J. S. L. Tsui. 2000. Discretionary accruals models and audit qualifications. *Journal of Accounting and Economics* 30: 241–252.

———, D. Givoly, and C. Hayn. 2002. The rewards to meeting or beating earnings expectations. *Journal of Accounting and Economics* 33: 173–204.

———, and P. Mohanran. 2004. Private information, earnings manipulations, and executive stock-option exercises. *The Accounting Review* 79 (October): 889–920.

Beasley, M. S., J. V. Carcello, and D. R. Hermanson. 1999. *Fraudulent Financial Reporting: 1987–1997, An Analysis of U.S. Public Companies.* Committee of Sponsoring Organizations of the Treadway Commission. New York, NY: COSO.

Becker, C. L., M. L. DeFond, J. Jiambalvo, and K. R. Subramanyam. 1998. The effect of audit quality on earnings management. *Contemporary Accounting Research* 15: 1–24.

Burgstahler, D., and I. Dichev. 1997. Earnings management to avoid earnings decreases and losses. *Journal of Accounting and Economics* 24: 99–126.

Butler, M., A. J. Leone, and M. Willenborg. 2004. An empirical analysis of auditor reporting and its association with abnormal accruals. *Journal of Accounting and Economics* 37 (June): 139–165.

Byrnes, N., R. A. Melcher, and D. Sparks. 1998. Earnings hocus-pocus: How companies come up with the numbers they want. *BusinessWeek* (October 5): 135.

Carcello, J., and T. L. Neal. 2000. Audit committee composition and auditor reporting. *The Accounting Review* 75 (October): 453–468.

Cheng, S. 2004. R&D expenditures and CEO compensation. *The Accounting Review* 79: 305–328.

Chung, H., and S. Kallapur. 2003. Client importance, non-audit fees, and abnormal accruals. *The Accounting Review* 78 (October): 931–955.

Clikeman, P. 1998. Auditor independence: Continuing controversy. *Ohio CPA Journal* 57 (2): 40–43.

Collingwood, H. 2001. The earnings game: Everyone plays, nobody wins. *Harvard Business Review* (June): 65–74.

DeAngelo, L. E. 1988. Managerial competition, information costs, and corporate governance: The use of accounting performance measures in proxy contests. *Journal of Accounting and Economics* 10: 3–36.

Dechow, P. M., and R. G. Sloan. 1991. Executive incentives and the horizon problem. *Journal of Accounting and Economics* 14: 51–89.

———, ———, and A. P. Sweeney. 1995. Detecting earnings management. *The Accounting Review* 70 (April): 193–225.

———, ———, and ———. 1996. Causes and consequences of earnings manipulation: An analysis of firms subject to enforcement actions by the SEC. *Contemporary Account Research* 13 (Spring): 1–36.

———, and I. D. Dichev. 2002. The quality of accruals and earnings: The role of accrual estimation errors. *The Accounting Review* 77 (Supplement): 35–59.

DeFond, M. L., and J. Jiambalvo. 1994. Debt covenant violation and manipulation of accruals. *Journal of Accounting and Economics* 17: 145–176.

———, and C. W. Park. 2001. The reversal of abnormal accruals and the market valuation of earnings surprises. *The Accounting Review* 76 (July): 375–404.

Denis, D. J., and D. K. Denis. 1995. Performance changes following top management dismissals. *Journal of Finance* 50: 1029–1057.

Dowdell, T. D., and J. Krishnan. 2004. Former audit firm personnel as CFOs: Effects on earnings management. *Canadian Accounting Perspectives* 3 (1): 117–142.

Francis, J. R., E. L. Maydew, and H. C. Sparks. 1999. The role of Big 6 auditors in the credible reporting of accruals. *Auditing: A Journal of Practice & Theory* 18 (Fall): 17–34.

Frankel, R. M., M. F. Johnson, and K. K. Nelson. 2002. The relation between auditors' fees for nonaudit services and earnings management. *The Accounting Review* 77 (Supplement): 71–105.

Geiger, M. A., and P. Taylor. 2003. CEO and CFO certifications of financial information. *Accounting Horizons* (December): 357–368.

———, D. S. North, and B. O'Connell. 2005. The auditor-to-client revolving door and earnings management. *Journal of Accounting, Auditing and Finance* 20 (1): 1–26.

Givoly, D., and C. Hayn. 2000. The changing time-series properties of earnings, cash flows and accruals: Has financial reporting become more conservative? *Journal of Accounting and Economics* 29: 287–320.

Healy, P. M., and J. M. Wahlen. 1999. A review of the earnings management literature and its implications for standard setting. *Accounting Horizons* 13 (December): 365–383.

Jones, J. 1991. Earnings management during import relief investigations. *Journal of Accounting Research* 29 (Autumn): 193–228.

Kinney, W. R., and R. D. Martin. 1994. Does auditing reduce bias in financial reporting? A Review of audit-related adjustment studies. *Auditing: A Journal of Practice & Theory* 13 (Spring): 37–48.

Klein, A. 2002. Audit committee, board of director characteristics, and earnings management. *Journal of Accounting and Economics* 29: 85–113.

Kothari, S. P., A. J. Leone, and C. E. Wasley. 2005. Performance matched discretionary accruals. *Journal of Accounting and Economics* 39: 163–197.

Kreutzfeldt, R. W., and W. A. Wallace. 1986. Error characteristics in audit populations: Their profile and relationship to environmental factors. *Auditing: A Journal of Practice & Theory* 6 (Fall): 20–43.

Lennox, C. 2005. Audit quality and executive officers' affiliations with CPA firms. *Journal of Accounting and Economics* 39: 201–231.

Levitt, A. 1998. The numbers game. Remarks delivered at NYU Center for Law and Business, New York, NY, September 28.

———. 2000. Renewing the covenant with investors. Remarks delivered at NYU Center for Law and Business, New York, NY, May 10.

Matsumoto, D. A. 2002. Management's incentives to avoid negative earnings surprises. *The Accounting Review* 77 (July): 483–514.

Menon, K., and D. D. Williams. 2004. Former audit partners and abnormal accruals. *The Accounting Review* 79 (October): 1095–1118.

Mian, S. 2001. On the choice and replacement of chief financial officers. *Journal of Financial Economics* 60: 143–175.

Murphy, K. J., and J. L. Zimmerman. 1993. Financial performance surrounding CEO turnover. *Journal of Accounting and Economics* 16: 273–316.

Nelson, M. W., J. A. Elliot, and R. L. Tarpley. 2002. Evidence from auditors about managers' and auditors' earnings management decisions. *The Accounting Review* 77 (Supplement): 175–202.

Parrino, R. 1997. CEO turnover and outside succession: A cross-sectional analysis. *Journal of Financial Economics* 46: 165–197.

Pincus, M., and S. Rajgopal. 2002. The interaction between accrual management and hedging: Evidence from oil and gas companies. *The Accounting Review* 77 (January): 127–160.

Pourciau, S. 1993. Earnings management and nonroutine executive changes. *Journal of Accounting and Economics* 16: 317–336.

Reynolds, J. K., and J. R. Francis. 2000. Does size matter? The influence of large clients on office-level auditor-reporting decisions. *Journal of Accounting and Economics* 30 (December): 375–400.

Securities and Exchange Commission (SEC). 2000. *Final Rule: Revision of the Commission's Auditor Independence Requirements.* File No. S7-3-00. Washington, D.C.: SEC.

———. 2002. *Waste Management Founder, Five Other Former Top Officers Sued for Massive Fraud.* Press Release 2002-44. Office of the Division of Enforcement. March 26. Washington, D.C.: SEC.

Schipper, K. 1989. Commentary on earnings management. *Accounting Horizons* 3 (4): 91–102.

Smith, E. B. 2004. What puts the CFO in the middle of scandals? *USA Today* (March 4).

Strong, J. S., and J. R. Meyer. 1987. Asset writedowns: Managerial incentives and security returns. *Journal of Finance* 42: 643–661.

Teoh, S., I. Welch, and T. Wong. 1998. Earnings management and the long-run market performance of initial public offerings. *Journal of Finance* 53: 1935–1974.

U.S. House of Representatives. 2002. The Sarbanes-Oxley Act of 2002. Public Law 107-204 [H. R. 3763]. Washington, D.C.: Government Printing Office.

U.S. Senate. 2002. Hearings before the subcommittee on telecommunications and finance of the committee on energy and commerce. February 2. Washington, D.C.: Government Printing Office.

Xie, B., W. N. Davidson III, and P. J. DaDalt. 2003. Earnings management and corporate governance: The role of the board and audit committee. *Journal of Corporate Finance* 9 (June): 295–316.

Yermack, D. 1997. Good timing: CEO stock option awards and company news announcements. *Journal of Finance* 52: 449–476.

Zmijewski, M. 1984. Methodological issues related to the estimation of financial distress. *Journal of Accounting Research* 22 (Supplement): 59–82.

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A Rational Expectations Theory of Kinks in Financial Reporting

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ABSTRACT: We present a rational model of earnings management. An informed manager, whose compensation is linked to the stock price, trades off the benefit of boosting the stock price by inflating the reported earnings against the costs of such manipulation. The investors rationally interpret his actions and adjust the price accordingly. When the distribution of true earnings and the compensation scheme are smooth, the conventional equilibrium in this signaling framework is also smooth and fully revealing. In this paper, we show that in the same "smooth" environment there exist equilibria in which kinks and discontinuities emerge endogenously in the distribution of reported earnings. The manager optimally chooses a partially pooling strategy, introducing endogenous noise into his report. The resulting vagueness enables the manager to reduce the average manipulation costs. The equilibrium has perfect revelation of earnings in the right and left tails of the distribution, while for intermediate earnings realizations, we get one or more pools that manifest themselves as discontinuities in the distribution of reported earnings. We study the properties of these partially pooling equilibria and suggest applications to financial reporting.

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I. INTRODUCTION

Managers of publicly traded firms possess information that is crucial for valuation of their firms. Lacking access to their private information, investors must determine the stock price primarily based on public information. Consequently, mandatory periodical earnings announcements by managers serve as the predominant source of information for investors. Managers have some leeway in reporting earnings within the accounting conventions. Moreover, the managerial compensation is frequently related to earnings either directly or indirectly via the stock price. Thus, managers have the ability and the incentives to manage (manipulate) earnings, and the empirical evidence suggests that they frequently do so.[1] There is also evidence that managerial compensation schemes affect the degree of earnings management.[2]

Theoretical papers on earnings management have focused on equilibria in which the distribution of reported earnings is smooth. One strand of the literature develops separating reporting equilibria in which investors can back out the private information of the manager from his report (e.g., Riley 1979; Stein 1989). Another approach (Fischer and Verrecchia 2000) is to introduce exogenous noise that precludes the investors from fully recovering the information of the manager out of his report. Regardless of the exact modeling approach, the result is a smooth distribution of reported earnings with no kinks or discontinuities. This prediction seems at odds with the extant evidence that identifies discontinuities in the reported earnings distribution.

We suggest a rational model in which kinks in reported earnings emerge endogenously, even though both the distribution of true earnings and the compensation of the manager are smooth. We develop the model in the context of earnings management, but suggest that several other financial reporting environments fit into this basic framework.

We use a signaling model of earnings management similar to Fischer and Verrecchia (2000), where an informed manager trades off the costs of earnings management against the benefit of higher stock price resulting from higher reported earnings. Investors rationally interpret the reported earnings in formulating their beliefs and set the price accordingly. This trade-off determines the manager's optimal level of manipulation for each level of realized earnings.

Truthful reporting by the manager is not an equilibrium, since if investors believe the report, then managers have an incentive to inflate the reported earnings. A standard equilibrium in this setting is a perfectly separating one: each manager "type" inflates his report,[3] however, investors correctly anticipate this action and infer the true earnings. Still, the manager incurs the cost of earnings management. The perfectly separating equilibrium creates a continuous distribution of reports and, hence, no kinks arise. In this paper, we show that a multitude of other equilibria exist in this setting, where reported earnings discontinuities arise endogenously.

We first claim that the following is a Perfect Bayesian Equilibrium: If the earnings are either low or high, then the manager issues a report that reveals his type, as in the perfectly separating equilibrium.[4] However, if the manager observes earnings in an intermediate range, then he always reports the same amount (a pooling report); i.e., the investors can infer only the interval of types to which he belongs, but not his actual type. The pooling behavior of the manager creates an endogenous discontinuity in the distribution of reports. This partially pooling equilibrium is supported by rational investors' beliefs.

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[1] The empirical literature on earnings management is extensive. See Healy and Wahlen (1999) for a review.

[2] See, for instance, Healy (1985), Bergstresser and Philippon (2006), Burns and Kedia (2006), and Cheng and Warfield (2005).

[3] We sometimes refer to the realization of true earnings as the manager's "type" to be consistent with the signaling literature.

[4] The manager does not report the truth, but his report does reveal the true earnings to the investors, who then price the stock correctly.

We refer to the range of the earnings realizations that result in the same report as the "pooling interval." We show that both the size (the length) of the pooling interval and its location relative to the mean of the distribution are determined endogenously. The size of the pooling interval increases in the sensitivity of the manager's payoff to the stock price and decreases in the unit cost of earnings manipulation. The location of the pool is not only affected by these parameters, but also by the volatility of earnings distribution—as the volatility increases, the interval moves away from the mean into the left tail.

We show that the manager strictly prefers this partially pooling equilibrium over the standard separating equilibrium, since it allows him to reduce the costs of earnings management by hiding some of his private information in a pooled report. Unlike Fischer and Verrecchia (2000), where the noise is exogenous and stems from the market uncertainty regarding managerial incentives, our analysis suggests that noise can be an endogenous phenomenon.

The separating equilibrium imposes no risk on the investors since they can always back out the information from the report and price the stock precisely. The partially pooling equilibrium does not always allow the investors to infer the manager's information precisely; hence, they face more risk. Consequently, whether investors prefer the partially pooling equilibrium or the separating one depends on their risk preferences, as well as on the compensation they receive for bearing this risk, as reflected in the risk premium.

If investors are risk-neutral, then they are indifferent between the two equilibria, since in both cases the stock is priced correctly on average. If investors are risk-averse, then the result is more subtle. To illustrate it we discuss two cases: In the first case, investors have mean-variance preferences, and the market price is set so they hold a fixed supply of shares. In this case, perhaps surprisingly, the investors prefer the partially pooling equilibrium, although it imposes more risk. The additional risk is factored into the market price, and the risk premium more than compensates the investors. In the second case, investors minimize a quadratic loss function, where risk is not factored into the market price. In this case, investors prefer the less risky separating equilibrium.

We next turn to characterize general properties of all possible equilibria in this model. We show that in any equilibrium the report must be a non-decreasing function of true earnings, and that any discontinuity must be associated with a pooling interval similar to the one discussed above. We also demonstrate the existence of partially pooling equilibria with multiple pooling intervals. All these equilibria involve a separating range on the left-hand side and right-hand side of the type distribution, and multiple (as many as we like) pooling intervals in the middle.[5] The basic partially pooling equilibrium with a single pooling interval serves as a building block in constructing the more general equilibria.

All the equilibria we identify consist of areas of the standard separating equilibrium and one or more pooling intervals. A special case is the standard separating equilibrium on both tails of the distribution combined with a step function reporting strategy, similar to those obtained in the "cheap talk" literature (see Crawford and Sobel 1982). In "cheap talk" models, misreporting is costless and so perfect revelation is impossible. The only possible equilibria in those cases are "babbling" equilibria, in which no information is conveyed, and step-function equilibria, which are partially revealing but do not reveal the full extent of private information on any interval of types. By contrast, in our costly signaling model, the only possible equilibria are either fully revealing or partially revealing ones. "Babbling" is not an equilibrium in our setting.

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[5] We can also construct equilibria where there are pooling intervals buffered by separating intervals.

Our results can be applied to a wide range of settings in financial reporting. First, in the context of earnings management, it has been documented that the empirical distribution of accounting earnings is discontinuous. Burgstahler and Dichev (1997), Degeorge et al. (1999), and Abarbanell and Lehavy (2003) provide evidence on such discontinuities, which manifest themselves as kinks in the distribution of accounting earnings.[6] Such kinks are found slightly below certain predetermined earnings levels—zero earnings, last year's earnings (zero change), and analysts' consensus expectation. The evidence suggests that managers manipulate earnings as if to meet exogenously pre-specified targets, such as avoiding losses, avoiding a decrease in accounting earnings, or meeting analysts' forecasts.[7]

A possible source for these kinks could be a discontinuity in the managerial compensation. For instance, a bonus paid upon reaching a pre-specified accounting earnings target would cause all managers with true earnings slightly below the target to report earnings meeting the target—as long as the cost of earnings manipulation is lower than the bonus. However, this explanation appears to be incomplete for two main reasons. First, there is little evidence that managers are paid bonuses explicitly for meeting analysts' expectations or for exceeding zero earnings. Second, the "bonus" explanation does not apply to stock-based compensation, since in order to reach a specific stock price target, managers must manage not only the earnings, but also the expectations and beliefs of investors who set the price. Thus, kinks and discontinuities in the distribution of accounting earnings are likely to be also driven by forces beyond the obvious ones, especially when stock-based compensation plays a dominant role.[8] Our theory suggests an alternative source for kinks in the distribution of earnings. Self-fulfilling market expectations accompanied by a pooling behavior by managers, whose compensation is tied to the stock price, manifest themselves as endogenous kinks in the distribution of reported earnings.

Our results of the pooling behavior for intermediate values combined with a separating behavior at the tails can also be applied to other phenomena in financial reporting, in which an uninformed decision maker obtains information from an informed party. We suggest that internal financial reporting within a firm, internal capital allocation processes, the process of raising outside equity, as well as analysts' forecasts and recommendations, can all fit into the basic framework of this model. In all these cases, an informed expert issues a report to a less informed decision maker. The expert may choose a partially pooling reporting strategy to save on the manipulation costs. Extremely good and extremely bad recommendations are revealed, while intermediate outcomes are pooled together in a noisy report.

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[6] A discontinuity or a kink in the distribution of reported earnings is typically defined as an interval of earnings reports such that the frequency of the reports inside this interval is substantially lower than the frequency of the reports in similar intervals to the left and to the right of that interval (see, for example, Burgstahler and Dichev 1997).

[7] Some recent papers question the connection between the kink phenomenon (especially near zero) and discretionary earnings management. See Dechow et al. (2003), Beaver et al. (2004), and Durtschi and Easton (2005).

[8] A behavioral explanation for kinks in the distribution of reported earnings was suggested by Dye (2002). The idea is that investors cannot fully comprehend the magnitude of earnings, and therefore classify firms into two categories such as "loss" firms and "profit" firms. Furthermore, investors assign favorable attributes only to the latter category. This motivates managers to engage in "classifications management," yielding a kink at 0.

Our model adds to a group of *one-period* models studying earnings management. These models are surveyed in Lambert (2001), and include papers such as Verrecchia (1986), Baiman et al. (1987), Newman and Sansing (1993), Evans and Sridhar (1996), and Demski and Dye (1999). Most relevant to our paper is Fischer and Verrecchia (2000) who offer a one-period model of earnings bias. Studying earnings management in a one-period framework ignores intertemporal aspects of earnings management such as earnings smoothing, dynamic rebalancing of discretionary accruals, and reputation building. We do not address those issues in this paper.[9]

The paper is organized as follows. In Section II we present the model and the standard separating equilibrium. In Section III we derive the partially pooling equilibrium with one pooling interval, and study its properties. In Section IV we study the preferences of the manager and the investors over the different equilibria. In Section V we specify general properties of equilibria in the model, and show the existence of equilibria with multiple pooling intervals. Section VI discusses possible applications of the partially pooling equilibrium to different environments of financial reporting, while Section VII concludes. Technical proofs are in the Appendix.

II. MODEL

We assume that the true earnings of the firm, x, are drawn from a normal distribution with mean x_0 and variance σ^2. The cumulative distribution is denoted by F, and the density is denoted by f. The parameters of the distribution are common knowledge; however, only the manager observes the realization x.[10] The manager is mandated to issue an earnings report, x^R, which the investors observe and use to price the stock. The payoff of a manager who reports x^R after observing x is given by:

$$U^M(x, x^R) = \alpha P(x^R) - \beta(x - x^R)^2, \tag{1}$$

where $\alpha \geq 0$, $\beta > 0$, and $P(x^R)$ is the market price of the firm given the report.[11]

The first term of the manager's payoff function represents the fact that managerial compensation depends on the stock price. We refer to α as the sensitivity of the manager's payoff to stock price. The second term of the manager's payoff function represents the cost of manipulation. Positive β implies that higher deviations from the truth carry higher legal, regulatory, or psychic costs for the manager. An alternative interpretation of this term is that the manager can take real actions (e.g., asset sales, suboptimal investments, aggressive sales efforts) to generate earnings that are different from the earnings under the optimal policy. These actions are costly in terms of the future earnings of the firm and, therefore, are costly for the manager.[12]

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[9] Dye (1988) offers both one- and two-period models of earnings management. He suggests that shareholders may not eliminate the tendency of managers to engage in earnings management. Earnings smoothing has received a lot of theoretical attention. See for instance Gonedes (1972), Barnea et al. (1975), Lambert (1984), Dye (1988), Trueman and Titman (1988), Fudenberg and Tirole (1995), Goel and Thakor (2003), and the survey by Ronen and Sadan (1981). These papers do not show how kinks can be generated in the distribution of accounting earnings.

[10] While x denotes the realization of earnings known only to the manager, we denote by $\tilde{x}$ the random variable of earnings observed by the uninformed investors.

[11] All parameters are common knowledge. Technically the model works for $\alpha < 0$ as well with appropriate changes to the comparative statics and to the results regarding the location of the pooling interval.

[12] See for example Ewert and Wagenhofer (2005) and Stein (1989).

We assume that investors are risk-neutral and, hence, price the stock proportional to the expected value conditional on all the available information.[13] Thus, the price of the stock prior to the manager's report is x_0—the prior mean, while the post-report price is $P(x^R) = E(\tilde{x}|x^R)$.[14] Plugging into (1) yields:

$$U^M(x, x^R) = \alpha E(\tilde{x}|x^R) - \beta(x - x^R)^2. \tag{2}$$

The manager has two conflicting interests: on the one hand, he would like to boost the stock price by manipulating his report; on the other hand, he does not want to manipulate his report too much, because the marginal cost of manipulation is increasing.[15] The relative weight that the manager assigns to each of these two incentives is determined by the ratio $\frac{\alpha}{\beta}$. The higher this ratio is, the more inclined is the manager to deviate from the truth. The combination of the normal distribution of earnings with a quadratic cost function yields a tractable model.

A reporting strategy for the manager is a real function $\rho: R \rightarrow R$ that maps true earnings into reports: $x^R = \rho(x)$.[16] A pricing function for the investors is a function $P: R \rightarrow R$ that maps the manager's report into a price. A "Perfect Bayesian Equilibrium" is defined as a reporting strategy ρ^* for the manager, together with a pricing function P^* for the investors, such that:[17]

1. The pricing function P^* is consistent with the strategy ρ^*, by applying Bayes rule whenever possible.
2. For all $x \in R$, $\rho^*(x) \in \arg\max_{x^R} U^M(x, x^R)$.

Observe first that truthful reporting, i.e., $\rho(x) = x$, is not an equilibrium. Indeed, if $\rho(x) = x$ for all $x \in R$, then investors adjust their beliefs to reflect this strategy; thus, $P(x^R) = x^R$ for all reports x^R. If a manager who observes real earnings of x reports truthfully, then he obtains αx. If this manager raises his report to $x + \varepsilon$ ($\varepsilon > 0$), then he obtains $\alpha(x + \varepsilon) - \beta\varepsilon^2$. Thus, deviation is beneficial as long as $0 < \varepsilon < \frac{\alpha}{\beta}$. The same argument shows that truth telling cannot be an equilibrium over any interval of a positive length.

The conventional equilibrium in this model is the perfectly separating (fully revealing) one. This equilibrium serves as a benchmark for our analysis.

Proposition 1: There exists a unique perfectly separating, continuously differentiable equilibrium. The equilibrium strategy of the manager is linear in x:

---

[13] This assumption is common in the literature, see for example, Fischer and Verrecchia (2000), Sankar and Subramanyam (2001), and Verrecchia (2001). In Section IV, we discuss the objective function of the investors in greater detail.

[14] In earlier versions of this paper, we assumed that $p_1 = cE(\tilde{x}|x^R)$, where $c > 0$ is the pricing coefficient for expected economic earnings. In what follows, we assume $c = 1$, as is common in one-period models. All of our results hold in the more general case.

[15] Note that the manager determines his report only after the true earnings are realized. Since the pricing function of the investors is known to the manager (and is not stochastic), the manager does not face any uncertainty. Thus, the risk attitude of the manager in this model is irrelevant.

[16] We restrict attention to pure strategies, i.e., given the manager's type x, he chooses a report $\rho(x)$, and does not randomize over several reports. In Section V we show that this restriction is essentially without loss of generality.

[17] As shown in Fudenberg and Tirole (1991), in our one-period model, any Perfect Bayesian Equilibrium is also a "Sequential Equilibrium" (Kreps and Wilson 1982).

$\rho_s^*(x) = x + \frac{\alpha}{2\beta}$ for all $x \in R$. The pricing function is linear in the report: $P_s^*(x^R) = x^R - \frac{\alpha}{2\beta}$ for all $x^R \in R$.

This perfectly separating equilibrium is characterized by a constant level of earnings manipulation. Regardless of the realization of true earnings, the manager inflates his report by $\frac{\alpha}{2\beta}$. Naturally, the investors are not fooled and price the stock correctly.

In deriving this equilibrium we used just one feature of the normal distribution: The fact that its support is the entire real line. In fact, this separating equilibrium applies to any continuous distribution that has a positive density on the entire real line.[18]

This type of separating equilibrium is standard in the continuous-type, costly signaling literature, such as Riley (1979). Notice that this equilibrium is inefficient. The manager cannot avoid the costly earnings management, and pays the costs that it imposes on him in the amount of $\frac{\alpha^2}{4\beta}$. Unfortunately, he gains nothing from this behavior, because it is correctly interpreted by the investors.[19]

III. A PARTIALLY POOLING EQUILIBRIUM

The separating equilibrium yields a smooth distribution of reported earnings and, hence, no kinks. However, there exist other equilibria in this model. These equilibria are partially pooling in the sense that they are separating for some earnings realizations, while pooling for others. Thus, for a range of earnings realizations, all managers will issue the same report (a pool). This behavior creates kinks in the distribution of reported earnings, although both the earnings distribution and the compensation are smooth.

In this section we demonstrate the simplest of these equilibria, which involves just one pooling interval where all the types in the pooling interval report the highest value of the interval. This equilibrium generates a single kink in the distribution of reported earnings. In Section V, we generalize our construction to show the existence of other equilibria with one pooling interval, as well as equilibria with multiple pooling intervals, leading to multiple kinks.

The equilibrium presented in this section provides the simplest illustration for the nature of pools and the resulting kink in reported earnings; it illustrates the benefits of pooling in terms of the savings in manipulation cost, and it serves as the primary building block for the other equilibria discussed in Section V. All the comparative statics results and the welfare analysis of this simple partially pooling equilibrium go through for the other equilibria developed in Section V.

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[18] A separating equilibrium of this type exists also for distributions with a bounded support, but in that case it cannot be given in an explicit form. Kadan and Yang (2005) use such a separating equilibrium to study the earnings management properties of option-based compensation.

[19] The empirical evidence on whether investors are actually fooled by earnings management is mixed. Rangan (1998) and Teoh et al. (1998) claim that managers succeed in fooling investors by manipulating reports. On the contrary, Shivakumar (2000) concludes that investors are not misled and account correctly for the manipulative behavior of managers.

Existence of a Partially Pooling Equilibrium

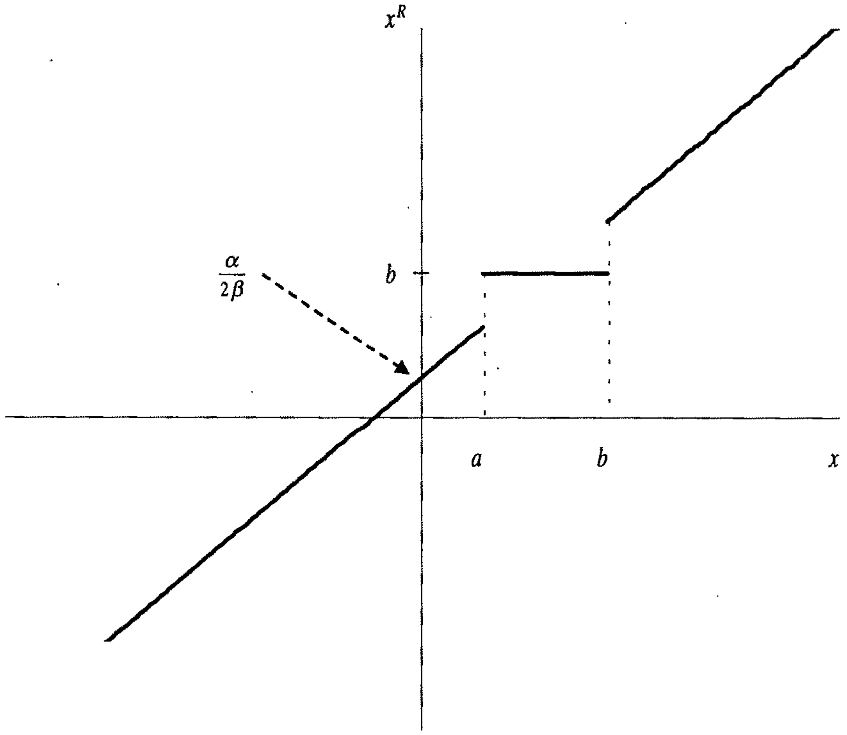
The idea is to transform the fully separating equilibrium in Proposition 1 into a partially pooling equilibrium. To establish this, we conjecture the existence of an interval $[a,b]$ such that the following partially pooling strategy is optimal for the manager:

$$\rho_p^*(x) = \begin{cases} b & x \in [a,b] \\ x + \dfrac{\alpha}{2\beta} & \text{otherwise} \end{cases}. \tag{3}$$

The conjectured strategy $\rho_p^*(x)$ is a simple modification of the fully revealing equilibrium strategy $\rho_s^*(x)$. For high or low values of earnings outside the interval $[a,b]$, the manager sticks to the same strategy as in Proposition 1: $\rho_p^*(x) = \rho_s^*(x) = x + \frac{\alpha}{2\beta}$. He does not report truthfully (this is suboptimal), but does reveal his true type. For the intermediate values that fall inside the interval $[a,b]$, the manager always reports b, the upper bound of the interval. Figure 1 illustrates this partially pooling strategy.

We would like to show that there exists an equilibrium based on such a partially pooling reporting strategy. The first step in proving existence is to find necessary conditions for $\rho_p^*(\cdot)$ to be an equilibrium. First, note that if $\rho_p^*(\cdot)$ is an equilibrium, then the type is fully

FIGURE 1
The Reporting Function in the Partially Pooling Equilibrium.



revealed for all reports $x^R < a + \frac{\alpha}{2\beta}$, or $x^R > b + \frac{\alpha}{2\beta}$. In contrast, when the investors observe a report of b, they can only deduce that the type is somewhere in the interval $[a,b]$. Using Bayes rule, it follows that conditional on a report of b, the posterior beliefs of the investors are distributed according to a truncated normal distribution over $[a,b]$.[20] Thus, the first necessary condition for a partially pooling equilibrium is that the pricing function of the investors must satisfy:

$$P_p^*(x^R) = \begin{cases} x^R - \frac{\alpha}{2\beta} & x^R < a + \frac{\alpha}{2\beta} \text{ or } x^R > b + \frac{\alpha}{2\beta} \\ d(a,b) & x^R = b, \end{cases} \tag{4}$$

where $d(a,b) \equiv E(\tilde{x}|\tilde{x} \in [a,b])$ is the mean of true earnings conditional on the information that they are in $[a,b]$. Next, notice that in order for $\rho_p^*(\cdot)$ to be an equilibrium, a manager with type $x = a$ must be indifferent between reporting $a + \frac{\alpha}{2\beta}$ and reporting b. Similarly, the manager with type $x = b$ must be indifferent between reporting $b + \frac{\alpha}{2\beta}$ and reporting b. Evaluating the manager's payoff given by Equation (2) at these points, and using Equation (4) we obtain a system of equations:

$$\begin{aligned} \alpha a - \beta\left(\frac{\alpha}{2\beta}\right)^2 &= \alpha d(a,b) - \beta(b-a)^2 \\ \alpha b - \beta\left(\frac{\alpha}{2\beta}\right)^2 &= \alpha d(a,b). \end{aligned} \tag{5}$$

Solving it yields:

$$b = a + \frac{\alpha}{\beta}, \tag{6}$$

and:

$$d(a,b) \equiv E(\tilde{x}|\tilde{x} \in [a,b]) = a + \frac{3\alpha}{4\beta}. \tag{7}$$

The intuition behind these necessary conditions is as follows. Both the "a" and the "b" types must be indifferent between the two alternatives they face. While the "a" type increases his earnings manipulation from $\frac{\alpha}{2\beta}$ to $\frac{\alpha}{\beta}$ $\left(\text{he reports } b = a + \frac{\alpha}{\beta} \text{ instead of } a + \frac{\alpha}{2\beta}\right)$, the "$b$" type gains by reducing his earnings manipulation costs to zero (he reports truthfully, instead of reporting $b + \frac{\alpha}{2\beta}\Big)$. Thus, the increase in earnings manipulation by the "*a*" type is identical to the decrease in earnings manipulation by the "*b*" type. However, because the manipulation cost is *convex*, the compensation in terms of price required by the "*a*" type exceeds the price concession made by the "*b*" type. This implies that *d(a,b)* lies to the right of the midpoint of the interval [*a*,*b*]. Given the quadratic cost function assumption, this conditional expectation lies exactly three quarters of the way between *a* and *b* (see Equation (7)).

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[20] This means that for all $s \in [a,b]$, $\Pr(\tilde{x} \leq s|\pi^R = b) = \Pr(\tilde{x} \leq s|\tilde{x} \in [a,b]) = \frac{F(s) - F(a)}{F(b) - F(a)}$.

The pricing function in Equation (4) is formed using Bayes rule given the conjectured equilibrium $\rho_p^*(\cdot)$. However, some potential reports never appear in this equilibrium. Specifically, no type ever issues a report that lies in $\left[a + \frac{\alpha}{2\beta}, b\right) \cup \left(b, b + \frac{\alpha}{2\beta}\right]$. Thus, to fully characterize the partially pooling equilibrium, we have to specify the out-of-equilibrium pricing. Since Bayes rule does not apply, we have some leeway in this choice. For simplicity, we use the following off-equilibrium beliefs: For all reports $x^R \in \left[a + \frac{\alpha}{2\beta}, b\right) \cup \left(b, b + \frac{\alpha}{2\beta}\right]$, the price is $P(x^R) = x^R - \frac{\alpha}{2\beta}$. Thus, if investors observe an unexpected report, they conclude that the manager is "mistakenly" playing the benchmark separating equilibrium $\rho_s^*(\cdot)$. Actually, there exists a continuum of out-of-equilibrium pricing functions (beliefs) that support the partially pooling equilibrium. Moreover, the out-of-equilibrium pricing functions can be made monotonically increasing in the reported earnings.[21]

Figure 2 depicts the pricing function of the investors in the partially pooling equilibrium. The dotted line describes the out-of-equilibrium pricing, while the bold dot is *d*(*a*,*b*): the price conditional on observing a report of $b = a + \frac{\alpha}{\beta}$.

Being necessary for the existence of the pool, Conditions (6) and (7) suggest that existence of a partially pooling equilibrium is not obvious. Does there always exist an interval [*a*,*b*], such that $b - a = \frac{\alpha}{\beta}$, while the conditional expectation over this interval lies at three quarters of the distance between *a* and *b*? Moreover, is the existence of such an interval sufficient for an equilibrium? In the next proposition we prove that the answers to these questions are positive, and so establish the existence of a partially pooling equilibrium.

Proposition 2: There exists a unique interval [*a*,*b*] such that the reporting strategy:

$$\rho_p^*(x) \equiv \begin{cases} b & x \in [a,b] \\ x + \frac{\alpha}{2\beta} & otherwise \end{cases}$$

joint with the pricing function:

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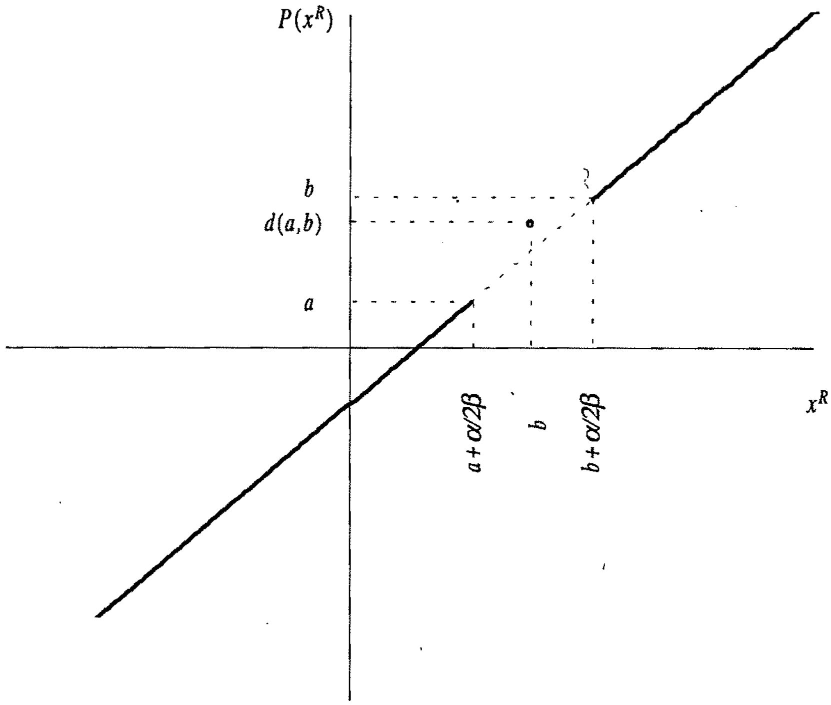
[21] A formal characterization of the set of off-equilibrium pricing functions that support the partially pooling equilibrium is somewhat technical. It is available upon request.

FIGURE 2
The Pricing Function in the Partially Pooling Equilibrium



$$P_p^*(x^R) = \begin{cases} x^R - \dfrac{\alpha}{2\beta} & x^R < a + \dfrac{\alpha}{2\beta} \text{ or } x^R > b + \dfrac{\alpha}{2\beta} \\ d(a,b) = a + \dfrac{3\alpha}{4\beta} & x^R = b \\ x^R - \dfrac{\alpha}{2\beta} & x^R \in \left[a + \dfrac{\alpha}{2\beta}, b\right) \cup \left(b, b + \dfrac{\alpha}{2\beta}\right] \end{cases}$$

constitute a Perfect Bayesian Equilibrium.[22]

The proof of this proposition is in the Appendix. It consists of two parts. The first part shows that if an interval [*a*,*b*] exists such that the necessary indifference Conditions (6) and (7) are satisfied, then the reporting function and beliefs specified in the proposition indeed constitute an equilibrium. The proof of this part is a consequence of the Spence-Mirrlees single crossing property, which we explore more generally in Section V. This part of the proof does not use properties of the distribution function except the non-bounded support.

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[22] Note that uniqueness here refers to the interval [*a*,*b*] and not to the equilibrium. See Section V for the derivation of other partially pooling equilibria in this model.

The second part of the proof shows that under the normal distribution assumption, such an interval actually exists and is unique. This part of the proof uses specific properties of the normal distribution.

Properties of the Partially Pooling Equilibrium

Below we study several properties of the partially pooling equilibrium. We start by studying the size and the location of the pooling interval relative to the mean of the earnings distribution. Next, we study the probability mass of the pooling interval, which is a measure of how pronounced we expect the kink to be. Finally, to provide some intuition for how our single firm model translates into a cross-sectional sample, we present cross-sectional simulations of the partially pooling equilibrium.

Size of the Pooling Interval

The following corollary shows that the size of the pooling interval depends solely on the parameters of the manager's payoff function.

Corollary 1: The size of the pooling interval is: $b - a = \frac{\alpha}{\beta}$.

Proof: Follows immediately from Condition (6).

Why does the length of the pooling interval increase in $\frac{\alpha}{\beta}$? To see the intuition, consider the following simple example. Assume that $\beta = 1$ and α increases from 1 to 1.5.[23] Suppose further, that following the increase in α the length of the pooling interval remains 1, and denote the upper end of the new pooling interval by b. Consider a manager with type $x = b - 1$. By switching from the separating to the pooling strategy, this manager gains from the increased price but incurs the cost of higher earnings management. Prior to the change in α, the increase in the cost was 0.75, and the increase in price (both prior to and after the increase in α) is also 0.75. Since prior to the change in α the coefficients of the price and cost were equal ($\alpha = \beta = 1$), the manager was indifferent between the two strategies (as required in equilibrium). Following the increase in α, the increase in the payoff due to the higher stock price is now higher (1.5×0.75 instead of 1×0.75). Moreover, the cost of joining the pool and reporting b instead of the new separating report is lower ($\frac{7}{16}$ versus 0.75 prior to the change). Both forces work in the same direction; hence, following the increase in α, type $b - 1$ strictly prefers to play according to the pooling strategy. In order for the partially pooling equilibrium to exist, the size of the pooling interval must increase. This will increase the cost that low types incur from joining the pool, offsetting the increased utility from the higher price.

Location of the Pooling Interval Relative to x_0

The next corollary identifies the location of the pooling interval relative to the unconditional mean of the true earnings.

Corollary 2: The following holds:

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[23] The coefficients α and β enter the model only as the ratio $\frac{\alpha}{\beta}$. Hence, the effect of a decrease in β is identical to an increase in α.

1. The conditional mean satisfies: $d(a,b) = a + \frac{3\alpha}{4\beta} = b - \frac{\alpha}{4\beta}$.
2. The conditional mean lies to the left of the unconditional mean: $d(a,b) < x_0$.
3. The lower bound of the pooling interval, a, is always to the left of the unconditional mean x_0. Moreover, $a < x_0 - \frac{3\alpha}{4\beta}$.
4. The upper bound of the pooling interval, b, can be to the left or to the right of the unconditional mean x_0, but is always smaller than $x_0 + \frac{\alpha}{4\beta}$.

Proof: The fact that $d(a,b) = a + \frac{3\alpha}{4\beta}$ is a restatement of Equation (7). To see that $d(a,b) < x_0$, note that $d(a,b) = a + \frac{3\alpha}{4\beta}$. The conditional mean lies to the right of the mid-point of the interval: $d(a,b) > \frac{a+b}{2}$. Under the normal distribution (as well as under any unimodal symmetric distribution) this is possible only if the conditional mean lies strictly to the left of the unconditional mean. Finally, since $d(a,b) < x_0$, $a = d(a,b) - \frac{3\alpha}{4\beta}$, and $b = d(a,b) + \frac{\alpha}{4\beta}$, we obtain $a < x_0 - \frac{3\alpha}{4\beta}$ and $b < x_0 + \frac{\alpha}{4\beta}$.

The pooling interval is either completely to the left of the unconditional mean or straddles the unconditional mean. Figure 3 illustrates this result. It shows the pooling interval compared to the distribution of true earnings (not the reported earnings). The figure presents a case where the pooling interval straddles the unconditional mean. The pooling interval can also lie entirely to the left of x_0 depending on parameter values. Due to the pooling, the distribution of reported earnings is not normal, and it is shifted to the right compared to the earnings distribution.[24] The shift is equal to $\frac{\alpha}{2\beta}$ outside of the pooling interval, and in the pooling interval it is on average smaller than $\frac{\alpha}{2\beta}$. The fact that $d(a,b)$ lies to the left of the unconditional mean implies that the probability mass of the pooling interval $[a,b]$ is concentrated to the left of the unconditional mean x_0.

The Effect of σ on the Location of the Pooling Interval

In the fully separating equilibrium, a change in the volatility of earnings has no effect on the equilibrium. In particular, earnings management is equal to $\frac{\alpha}{2\beta}$ regardless of σ. In

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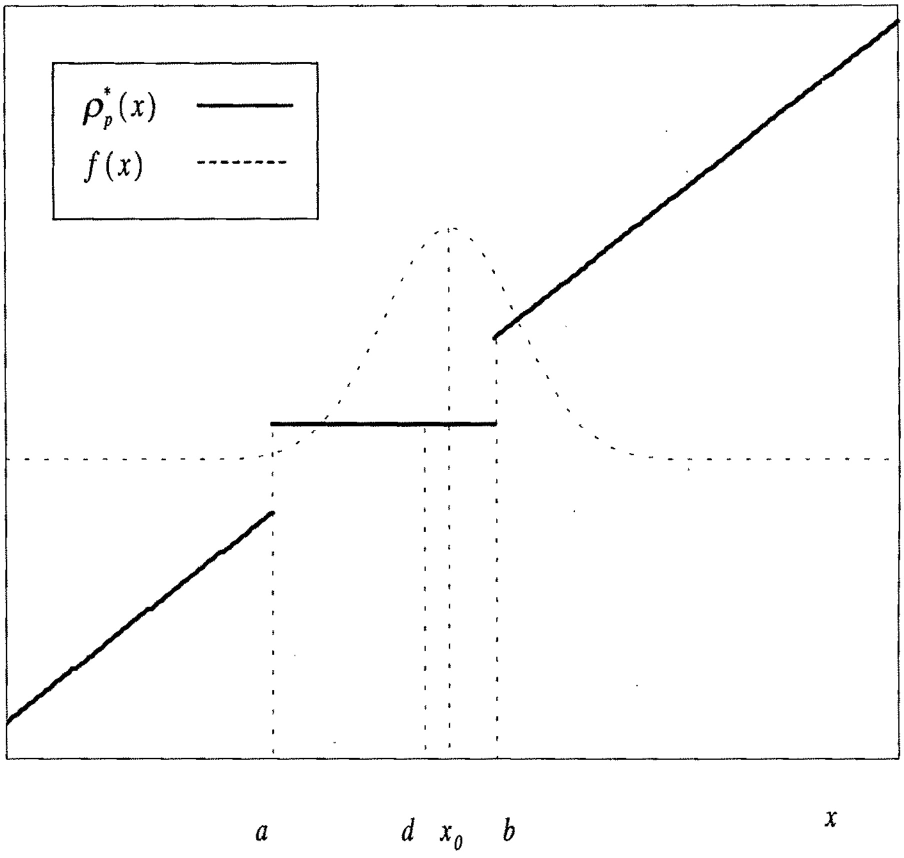
[24] The non-normality of the reports distribution induced us to assume a constant α. Fischer and Verrecchia (2000) assume that α is a normally distributed random variable to capture the uncertainty of investors regarding the value of this parameter. This is possible in their setting since they use a linear reporting function. If α were normally distributed in our setting, then the joint distribution of reports and α would be hard to track analytically, precluding the analysis of equilibrium.

FIGURE 3
The Partially Pooling Equilibrium and the Underlying Distribution



contrast, in the partially pooling equilibrium, a change in σ affects the location of the pooling interval $[a,b]$, and so affects the probability mass associated with the interval. To study the impact of σ on the location of the pooling interval relative to the mean of the earnings distribution, in the next corollary we focus on $a(\sigma)$, the left-hand side of the interval, and $d(\sigma)$, the conditional mean as functions of σ.

Corollary 3: A decrease in the underlying volatility moves the pooling interval and the conditional mean to the right: $\frac{\partial a(\sigma)}{\partial \sigma} < 0$ and $\frac{\partial d(\sigma)}{\partial \sigma} < 0$. Moreover: $\lim_{\sigma \to 0} a(\sigma) = x_0 - \frac{3\alpha}{4\beta}$, and $\lim_{\sigma \to \infty} a(\sigma) = -\infty$.

Corollary 3 shows that higher σ moves the pooling interval to the left of the distribution. On the other hand, as σ declines, the pooling interval moves to the right, eventually straddling the unconditional mean. In the limit, as σ tends to 0, the unconditional mean x_0 and the conditional mean $d(a,b) = a + \frac{3\alpha}{4\beta}$ coincide $\left(a = x_0 - \frac{3\alpha}{4\beta}, b = x_0 + \frac{\alpha}{4\beta}\right)$, and the probability of pooling goes to 1.

To understand the result, we must appeal to two properties of the normal distribution. First, the conditional expectation over an interval of length $\frac{\alpha}{\beta}$ moves closer to the upper bound of the interval, as the interval moves toward the left tail of the distribution. This can be seen as a result of the "thin tails" of the normal curve, implying that as one moves to the far left of the distribution, more relative weight is assigned to the right-hand side of the interval. Second, as one increases the variance of a normal distribution, the conditional expectation over a fixed interval that lies to the left of the unconditional mean must move to the left. Intuitively, this follows since higher variance implies a "flatter" normal curve, causing the conditional mean to be less tilted toward the right and the unconditional mean. Both of these properties are proven formally in the Appendix.[25]

Now, consider types $x = a$ and $x = b$, and think about how an increase in σ affects these two types. Before the increase in σ these two types were indifferent between the pooling and the separating strategies. If the pooling interval is unchanged, then from the second property above we know that $d(a,b)$ declines. But then the reward in price to type $x = a$ is not sufficient to make the manager indifferent between the separating and pooling strategies. Similarly, the price concession of type $x = b$ is too high, again breaking the indifference. To restore the indifference of the two types the location of the conditional mean must move to the right. Given the first property of the normal distribution mentioned above, this will happen if you shift the interval $[a,b]$ to the left, as required.[26]

The Probability of Pooling

Given the size and the location of the pooling interval, we can calculate the probability that the realized earnings fall into this interval. This probability is a measure of how pronounced we expect the kink to be in a cross-section of firms. As the probability of pooling goes to zero, the partially pooling equilibrium converges to the separating equilibrium. At the other extreme, when the probability of pooling goes to 1, the partially pooling equilibrium converges to a perfectly pooling equilibrium.

Consider the effect of σ on the probability of pooling. From Corollary 3 we know that as σ goes to infinity (all else being equal), the probability of pooling goes to zero. On the other hand, as σ goes to zero, the pooling interval straddles the unconditional mean and the probability of pooling goes to 1. This suggests that the probability of pooling decreases in σ. Namely, more volatile earnings induce less pronounced pooling. We are not able to prove this monotonicity analytically. The problem is that there is a region in the distribution in which the increase in variance has two conflicting effects (on the location of the interval and on the density function). However, extensive numerical calculations suggest this monotonicity. Neither could we show analytically that the probability of pooling increases monotonically in $\frac{\alpha}{\beta}$; however, numerical calculations suggest that this is indeed the case.

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[25] The first one is shown in the proof of Proposition 2, and the second in the proof of Corollary 3.

[26] The location of the pooling interval depends also on the ratio $\frac{\alpha}{\beta}$. We cannot provide an analytical solution of the effect of $\frac{\alpha}{\beta}$ on the location of the pooling interval. However, partial analytical results and numerical simulations suggest that as $\frac{\alpha}{\beta}$ increases, b increases monotonically, while a changes in a non-monotonic fashion. It increases for low values of $\frac{\alpha}{\beta}$ and, at a given point, starts to decrease monotonically.

Cross-Sectional Simulations

Note that for any given set of parameter values our model does not create a "dent" or a "kink" in the distribution of reports, but rather a "hole." In the partially pooling equilibrium, no report should be observed in the range $\left[a + \frac{\alpha}{2\beta}, b\right) \cup \left(b, b + \frac{\alpha}{2\beta}\right]$. This fact can be reconciled with the observed empirical results in a cross-section of firms. Each report takes place in a different environment, which implies that across these environments the holes in the distribution of reports do not overlap completely. Due to some overlap, however, we should observe a drop in the frequency of certain reports in the data, rather than a complete elimination of reports in an interval.

To illustrate this, Figure 4 presents the results of three separate simulations. Each one of the three graphs is a result of 10,000 rounds of simulation based on the partially pooling equilibrium. To take into account the variation in parameters across firms, we treat the mean, the standard deviation, and α as random variables in each round of simulation. Specifically, the mean x_0 is drawn from a normal distribution with mean 0 and standard deviation 0.1, and the standard deviation is drawn from a normal distribution (truncated at 0 to avoid negative variance) with mean 0.7 and standard deviation 0.1. The difference between the graphs stems from different values of $\frac{\alpha}{\beta}$. We assume that $\beta = 1$, and draw α from a normal distribution as follows: in the top graph α is drawn from a distribution with mean 1.8 and standard deviation 0.1, in the middle graph α is drawn from a distribution with mean 1.5 and standard deviation. 0.1, and in the bottom graph α is drawn from a distribution with mean 0.8 and standard deviation 0.1.

Consistent with the theoretical results above, the kink in the top graph is more pronounced than the kink in the middle graph. In the bottom graph, the values of $\frac{\alpha}{\beta}$ are sufficiently small, that the kink is virtually undetectable. We can generate a variety of kinks by changing the model's parameters.

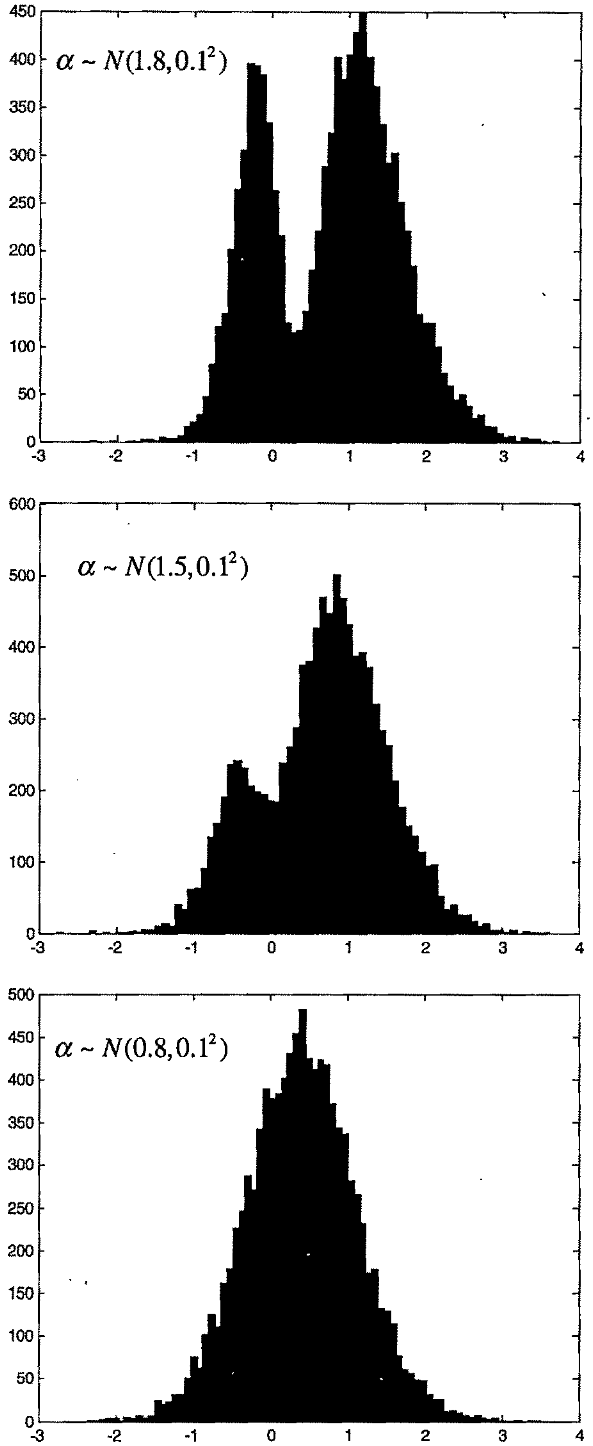
IV. PREFERRED EQUILIBRIUM

In this section we argue that managers always prefer the partially pooling equilibrium over the separating equilibrium. Investors' preferences depend on their assumed utility function. Risk-neutral investors are indifferent between the two equilibria, while risk-averse investors may prefer one equilibrium over the other, depending on the specific formulation of their utility function.

Equilibrium Preferred by the Manager

Our first observation is that the average earnings management in the partially pooling equilibrium is lower than in the separating equilibrium. By pooling all types in the interval $[a,b]$, the manager introduces endogenous noise (or vagueness) into his reports. This enables him to lower the average extent of earnings management.

FIGURE 4
Cross-Sectional Simulations



Proposition 3: The expected earnings management in the partially pooling equilibrium is lower than the expected earnings management in the separating equilibrium.[27]

In the empirical accounting literature, the extent of kinks in observed distribution of earnings reports is assumed to be an indicator of earnings management. However, from the above proposition, one can see that the expected earnings management in the partially pooling equilibrium is lower on average than in the separating one. Nevertheless, the pooling report is noisy, i.e., is characterized by a lower quality of information revelation, while the separating report can be perfectly deciphered by the investors.

As earnings management imposes a cost on the manager, one would expect that he would be *ex ante* better off in the partially pooling equilibrium, relative to the separating equilibrium. In fact, we can make a stronger statement. The manager prefers the partially pooling equilibrium not only in expectation (from an *ex ante* perspective), but also for each and every realization of earnings.

Proposition 4: For all $x \in (a,b)$, $U^M(x,\rho_p^*(x)) > U^M(x,\rho_s^*(x))$.

Intuitively, the pooling behavior of the manager has two effects on his payoff compared to the fully separating equilibrium. On the one hand, it changes the extent of the earnings management by the manager. On the other hand, it affects the pricing of the stock given this manipulation. The net effect is always positive within the pooling interval (a,b). If the realized earnings are such that the manager manipulates more in the pooling equilibrium, then he is more than compensated by the price increase relative to the separating equilibrium, while if he manipulates less, then the price decline is not sufficient to offset the reduction in cost.

An immediate implication of Proposition 4 is that the manager prefers the partially pooling equilibrium from an *ex ante* perspective as well, since he is better off for each and every realization. Again, the intuition here stems from the reduction of expected earnings management costs in the partially pooling equilibrium. Formally:

Corollary 4: The expected payoff of the manager is higher in the partially pooling equilibrium compared to the separating equilibrium.

Equilibrium Preferred by the Investors

The investors' ranking of the two equilibria depends on their objective function. Note that in the separating equilibrium there is no uncertainty left once the manager issues his report. In this equilibrium, the price always equals the actual realization of x. By contrast, in the partially pooling equilibrium, the price is equal to the realization on average, but not at each and every realization. Thus, investors' attitude toward risk matters.

In Section II we assumed that investors are risk-neutral and price the stock at its expectation conditional on the report. In this case, the two equilibria provide the investors with the same expected utility: they always break-even in the separating one, and break-even on average in the partially pooling one. Since the manager strictly prefers the partially pooling equilibrium we get Pareto dominance from the *ex ante* point of view.

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[27] It can also be easily shown that the expected cost of earnings management is lower in the partially pooling equilibrium than in the separating equilibrium.

Proposition 5: If investors are risk-neutral, then the partially pooling equilibrium *ex ante* Pareto dominates the separating equilibrium.

The dominance of the partially pooling equilibrium follows from the fact that the manager incurs lower expected manipulation costs. If investors would explicitly gain from the reports (e.g., by decreasing their uncertainty), then one would expect more of a horse race between the two equilibria.

The Pareto dominance result clearly relies on the risk-neutrality of the investors. How does risk aversion affect this result? We discuss below two cases of investors' risk aversion. In one case (mean variance preferences), investors prefer the partially pooling equilibrium despite their risk aversion, and in the other case (minimizing a quadratic loss function), they prefer the separating equilibrium.

Investors with Mean-Variance Preferences

Assume that investors maximize the certainty equivalent (*CE*) of their expected payoff with respect to the number of shares they want to hold at the beginning of the period, *n*. Assuming that the *CE* takes a standard mean-variance form, the optimization problem faced by investors is:

$$\max_{n} CE(x^R) \equiv (E(\tilde{x}|x^R) - P(x^R))n - 0.5zn^2Var(x|x^R), \tag{8}$$

where $z > 0$ is a proxy for the investors' coefficient of risk aversion. When the investors are risk-neutral, $z = 0$, the risk-neutrality of the investors implies that the only market-clearing price is, $P(x|x^R) = E(x|x^R)$ as assumed in the risk-neutral-investors case. Indeed, any price above this would lead to a "sellers only" market, whereas a price below the conditional expectation will lead to a "buyers only" market. One can think of this as being a subgame played after the manager issues his report; the outcome of this subgame is the market price.

The per-capita demand for stock is obtained by the first-order condition of Equation (8):

$$n^* = \frac{E(\tilde{x}|x^R) - P(x^R)}{zVar(x|x^R)}.$$

As usual in such models, the investor's demand increases in the expected value, but declines in the price of the stock, the coefficient of risk aversion, and the variance. Market-clearing condition requires that the per-capita demand must equal the exogenously given per-capita supply of the stock, denoted by $S(>0)$. This yields the equilibrium pricing function of the stock:

$$P^*_{MV}(x^R) = E(\tilde{x}|x^R) - SzVar(\tilde{x}|x^R), \tag{9}$$

where the subscript *MV* represents the model with mean-variance investors. Thus, the price for any given report must equal the conditional expectation less a risk premium that increases in the posterior variance.

Substituting the Price (9) into the certainty equivalent, and using the fact that $n^* = S$ (the market clears) we obtain:

$$CE(x^R = b) = S^2 zVar(\tilde{x}|\tilde{x} \in [a,b]) - 0.5zS^2 Var(\tilde{x}|\tilde{x} \in [a,b])$$
$$= 0.5S^2 zVar(\tilde{x}|\tilde{x} \in [a,b]). \qquad (10)$$

Thus, the certainty equivalent increases in the variance of investors' beliefs conditional on the report. This implies that if the same two types of equilibria exist in this case, then investors *ex ante* strictly prefer the partially pooling equilibrium. Note that in equilibrium, the pooling price $P^*_{MV}(b)$ will equal $a + \frac{3\alpha}{4\beta}$ (as in the risk-neutral case), but the conditional mean will be higher than $a + \frac{3\alpha}{4\beta}$ by the risk premium ($SzVar(\tilde{x}|x^R)$). Hence, from the manager's perspective, the equilibrium is qualitatively the same as in the risk-neutral case, and he strictly prefers it over the fully separating equilibrium. So, the partially pooling equilibrium *ex ante* Pareto dominates the separating equilibrium.[28]

This result seems surprising: The partially pooling equilibrium imposes additional risk on the risk-averse investors; how can they prefer this equilibrium? The answer lies in the fact that the additional risk is factored into the price in a way that more than compensates the investors. Higher variance makes the investors' demand for shares less elastic, and since the equilibrium holdings of all the investors remain the same as the market clears, the price must decline, which increases the consumer surplus of the investors.

Thus, risk aversion alone is not enough to make investors better off under the fully revealing equilibrium. Only when the price does not sufficiently compensate them for the additional risk they have to bear under the partially pooling equilibrium, will they strictly prefer the separating equilibrium. We illustrate such a setting next.

Investors Minimizing a Quadratic Loss Function

The second case we consider is when investors minimize the square of the difference between the price and the true earnings conditional on the report.[29] In this case, investors are risk-averse (they minimize a convex function) and yet the equilibrium stock price equals the expected earnings conditional on the report, just as in the case of risk-neutral investors. That is, the price reflects no risk premium. Since investors are not compensated for the risk they incur in the partially pooling equilibrium, they strictly prefer the separating equilibrium over the partially pooling one.

Obviously, this setup is quite extreme. In practice we expect that investors will demand a risk premium for the risk they face. Whether this risk premium is sufficient to induce them to prefer the partially pooling equilibrium over the separating one depends on the actual specification of the utility function.

Further Discussion

Which of the equilibria is more likely to prevail—the partially pooling or the fully separating? The game theoretic literature has no decisive answer to the equilibrium selection question. However, the following intuitive argument may provide a rationale for why the partially pooling equilibrium may prevail. The manager acts first; since he strictly prefers the partially pooling equilibrium (regardless of investors' utility function), he can announce prior to the realization of earnings that he is going to play according to the partially pooling equilibrium. This announcement is credible, since for each and every realization of earnings, he strictly prefers this equilibrium. The best response of the investors is then to comply and form beliefs based on this equilibrium strategy.

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[28] Existence of a partially pooling equilibrium in the mean-variance case is not obvious. Numerical calculations suggest that such an equilibrium always exists. Nevertheless, we have not been able to prove this analytically. Given this limitation, this welfare discussion applies only to cases where the partially pooling equilibrium exists.

[29] Guttman et al. (2006) consider such a case in the context of professional advice from biased experts.

Our model focuses on just one aspect of the stock-based compensation—the reporting part. In reality, stock-based compensation has a motivating aspect as well. Thus, an increase in α, while increasing earnings management, and the extent of pooling in our model, adds incentives for managers, and may induce a higher realization of true earnings. For this reason, our analysis in this section draws only a partial picture of the division of surplus between market participants.

V. GENERALIZATIONS OF THE PARTIALLY POOLING EQUILIBRIUM

In this section, we generalize the results obtained so far. First, we show that there exists a continuum of equilibria with a single pooling interval. Then, we show how these can be used as building blocks to construct partially pooling equilibria with multiple pooling intervals. All the equilibria we identify share similar properties with the partially pooling equilibrium obtained thus far. They all involve a separating equilibrium in the left-hand side and in the right-hand side of the distribution of earnings. In the middle we will have either one or several pooling intervals.

General Properties of Equilibria

The following proposition describes properties of any possible equilibrium in the model. All of the general properties below are not subject to the assumption that true earnings are normally distributed. To be completely general, we allow the manager of type x to randomize over several reports (i.e., use a mixed strategy).

Proposition 6: In any possible equilibrium:

1. The reporting function is non-decreasing in type.[30]
2. The number of discontinuities in the reporting function is either finite or countable and, hence, the set of types for which the reporting function is discontinuous has measure 0.
3. The set of types that randomize has measure 0.
4. For any equilibrium reporting function $\rho(\cdot)$, if $\rho(x)$ is discontinuous at $\hat{x}$, then there is a pooling interval either to the right or to the left of $\hat{x}$ (or both).

Thus, all the possible equilibria are non-decreasing and, hence, are continuous almost everywhere. The number of types for which the reporting function is discontinuous can be either finite or infinite, but the set of such types has measure (probability) 0. Moreover, Part 3 of the proposition shows that restricting attention to pure strategies is essentially nonbinding, since in any possible equilibrium, the number of types that randomize is negligible. Part 4 shows that pooling intervals are the only types of discontinuities that may occur in equilibrium.

The results so far enable us to get insights into the types of reporting functions that *cannot* constitute an equilibrium. We already know that reporting functions that decrease over an interval cannot exist in an equilibrium. Furthermore, the discussion in Section II shows that truth-telling cannot be part of an equilibrium reporting function on any interval. Finally, it is easy to see that "babbling" is not an equilibrium. By "babbling," we mean an equilibrium in which no information is conveyed.[31] This can be either due to the fact that all types issue the same report or that they all randomize over the same set of reports.

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[30] One has to be a little careful when using the term "non-decreasing" if the manager randomizes. Suppose that $x_1 < x_2$, and that type x_1 randomizes over reports in the set A_1 while type x_2 randomizes over reports in the set A_2. Then by "non-decreasing" we mean that $\inf A_2 \geq \sup A_1$. If the manager uses pure strategies, then this definition coincides with the standard notion of "non-decreasing."

Corollary 5: Babbling cannot be an equilibrium.

Other Equilibria with One Pooling Interval

In Section III we showed the existence of a partially pooling equilibrium, in which the manager reports the upper bound of the pooling interval. Given this report, the equilibrium pooling interval is determined uniquely. That equilibrium, however, is not the only partially pooling equilibrium in this model. Actually, there exists a continuum of similar partially pooling equilibria with a single pooling interval. Equilibria in this family differ by the pooling report that the manager makes when the realized earnings are in the pooling interval. While in our original equilibrium all managers with earnings in $[a,b]$ report the upper bound b, we can create other equilibria in which managers with earnings in the pooling interval choose a different report. Specifically, for every $\eta \in \left(-\frac{\alpha}{2\beta}, \frac{\alpha}{2\beta}\right)$ there exists a partially pooling equilibrium where all managers with earnings in the pooling interval $[a_\eta,b_\eta]$ report $b + \eta$, and for all other earnings they report according to the separating equilibrium. The size of the pooling interval, its location, and the mean of earnings conditional on being in the pooling interval are also determined by η. Our baseline equilibrium described in Section III corresponds to the case $\eta = 0$.

Proposition 7: Set $\eta \in \left(-\frac{\alpha}{2\beta}, \frac{\alpha}{2\beta}\right)$. There exists a unique interval $[a_\eta,b_\eta]$, where $b_\eta - a_\eta = \frac{\alpha}{\beta} - 2\eta$, such that the reporting strategy:

$$\rho_p^\eta(x) = \begin{cases} b_\eta + \eta & x \in [a_\eta,b_\eta] \\ x + \dfrac{\alpha}{2\beta} & otherwise \end{cases},$$

joint with the pricing function:

$$P^\eta(x^R) = \begin{cases} x^R - \dfrac{\alpha}{2\beta} & x^R < a_\eta + \dfrac{\alpha}{2\beta} \text{ or } x^R > b_\eta + \dfrac{\alpha}{2\beta} \\ b_\eta - \dfrac{\alpha}{4\beta} + \dfrac{\beta}{\alpha}\eta^2 & x^R = b_\eta + \eta \\ x^R - \dfrac{\alpha}{2\beta} & x^R \in \left[a_\eta + \dfrac{\alpha}{2\beta}, b_\eta + \eta\right) \cup \left(b_\eta + \eta, b_\eta + \dfrac{\alpha}{2\beta}\right] \end{cases}$$

constitute a Perfect Bayesian Equilibrium.

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[31] Babbling is a standard type of equilibrium in "cheap talk" models. See Crawford and Sobel (1982).

Proof: Similar to the proof of Proposition 2.

The family of equilibria described above nests our original partially pooling equilibrium as a special case. We have chosen to focus on this equilibrium first because it is the simplest one analytically, and completely represents the family. All the welfare implications, comparative statics, empirical predictions, and the robustness analysis carry on to this entire family of equilibria. The choice of η just changes the length and position of the pooling interval and the mean conditional on being in the pooling interval. In particular, in all these equilibria, the pooling interval is concentrated to the left of the unconditional mean.

It is worth mentioning that η has a monotonic effect on the size of the pooling interval, as well as on the location of the conditional mean relative to the pooling interval. However, η does not have a monotonic effect on welfare. To see this, note that when η increases, the size of the pooling interval decreases, but the pooling interval itself moves toward the unconditional mean. The first effect tends to decrease the probability mass of the pooling interval, while the second effect tends to increase it. As a result, the expected earnings management may increase or decrease following an increase in η. In the two extreme cases, $\eta \rightarrow -\frac{\alpha}{2\beta}$ and $\eta \rightarrow \frac{\alpha}{2\beta}$, the pooling equilibrium converges to the separating equilibrium. When $\eta \rightarrow -\frac{\alpha}{2\beta}$, the pooling interval moves toward the far left tail of the distribution and so its probability is negligible. In the case $\eta \rightarrow \frac{\alpha}{2\beta}$, the size of the pooling interval converges to 0, and so again its probability becomes negligible. Thus, the pooling behavior is most pronounced and, hence, the savings in earnings management to the manager are largest, for some intermediate value of η.

Equilibria with Multiple Pooling Intervals

Proposition 7 characterizes all the partially pooling equilibria in the model that have a single pooling interval. We show below how to construct other partially pooling equilibria with multiple pooling intervals by combining different η values in a single equilibrium. The key to such equilibria is the fact that for different η values, the location of the pooling interval (which is a function of η) is different. In fact, the left-hand bound of the pooling interval a_η can take any value below the unconditional mean x_0. Formally:

Lemma 1: Let $z < x_0$. There exists an $\eta \in \left(-\frac{\alpha}{2\beta}, \frac{\alpha}{2\beta}\right)$ such that $a_\eta = z$.

Note that as η goes to $\frac{\alpha}{2\beta}$, the size of the pooling interval $[a_\eta, b_\eta]$ (which is monotonic in η) goes to zero, and a_η tends to the mean of the prior distribution, x_0. On the other hand, as η goes to $-\frac{\alpha}{2\beta}$, the size of the pooling interval $[a_\eta, b_\eta]$ approaches $\frac{2\alpha}{\beta}$ (its maximal size), and a_η approaches $-\infty$.

Given Lemma 1, it is easy to construct equilibria with multiple pooling intervals. The idea is to choose η values such that the pooling intervals for these values do not overlap. This is always possible since the size of the pooling interval never exceeds $\frac{2\alpha}{\beta}$, while the support is infinite.

To flesh out the idea, consider the following example: Suppose you want to construct an equilibrium with two pooling intervals. Set $z_1 = x_0 - \frac{2\alpha}{\beta}$ and $z_2 = x_0 - \frac{4\alpha}{\beta}$. Then, by Lemma 1, you know that there exist η_1 and η_2 such that $a_{\eta_1} = z_1$ and $a_{\eta_2} = z_2$. The reporting function—which is separating on the entire real line except on the two pooling intervals $[a_{\eta_1}, b_{\eta_1}]$ and $[a_{\eta_2}, b_{\eta_2}]$—together with the pricing functions defined in Proposition 7 for each pooling interval constitute an equilibrium. Note that since $z_1 - z_2 = \frac{2\alpha}{\beta}$, the two pooling intervals do not overlap, because the size of pooling intervals is always smaller than $\frac{2\alpha}{\beta}$. This insures that the appropriate local indifference conditions for each pooling interval guarantee an equilibrium.

In general, one can use the same procedure to construct an equilibrium with any finite number of pooling intervals. For example, assume that you are interested in an equilibrium with N pooling intervals. Set $z_i = x_0 - i\frac{2\alpha}{\beta}$ for $i = 1,\dots,N$. For each i, Lemma 1 instructs that you can find an $\eta_i \in \left(-\frac{\alpha}{2\beta}, \frac{\alpha}{2\beta}\right)$ such that $a_{\eta_i} = z_i$. Moreover, the pooling intervals are defined by $[a_{\eta_i}, b_{\eta_i}]$, where $b_{\eta_i} - a_{\eta_i} = \frac{\alpha}{\beta} - 2\eta_i$ never overlap. Therefore, the indifference conditions for the individual a_{η_i} and b_{η_i} pairs insure that the reporting function accompanied by the appropriate pricing function is an equilibrium. We have thus shown:

Proposition 8: For any integer $N > 0$, there exists a partially pooling equilibrium with N pooling intervals.

Note that due to the way we constructed the equilibrium, any two pooling intervals are buffered by an interval of separating reports. But, this is not a necessity. By choosing appropriate η values, one can make the pooling intervals adjacent.[32] In this way, the reporting will take the form of a step function over a certain range. This is similar to the "cheap talk" partition equilibria identified in Crawford and Sobel (1982), but in a totally different setting of costly signaling.

The properties of all these partially pooling equilibria are very similar to those studied in the case of a single pooling interval. First, they all share the same qualitative properties. Pooling occurs for intermediate values of earnings realizations, and all the pools are concentrated to the left of the unconditional mean. Only the right-most pooling interval can (but does not have to) straddle the unconditional mean.

These equilibria also share the same welfare properties of the equilibrium with just one pooling interval. In all of them, the manager strictly prefers the pooling equilibrium over the separating equilibrium. Moreover, adding additional pooling intervals always reduces the expected earnings management.

VI. OTHER APPLICATIONS TO FINANCIAL REPORTING

Our analysis thus far has focused on the relation between managers and investors in the context of earnings management. Below we provide several examples of applications to financial reporting where partially pooling equilibria may be relevant.

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[32] Note that the number of pooling intervals can be either finite or infinite.

Internal Reporting

CEOs report to the board of directors, who must infer the true information about the company from the report. The compensation of the CEO clearly depends on the information provided to the board in addition to the mandatory public reports. Thus, the CEO has an incentive to boost the board's perception of her performance. The manager can manipulate the report itself, or make suboptimal operating and investment decisions designed to alter the reported performance. As before, the manager also faces costs associated with misrepresentations, including the possibility of dismissal. Contrary to external reports that are subject to audit and regulation, internal reports allow more leeway and managerial discretion. This suggests that the costs to misreporting internally are lower.

Our partially pooling equilibria suggest that the CEO is likely to choose the partially pooling equilibrium; extremely good and extremely bad outcomes are completely identified in equilibrium, while intermediate outcomes are pooled together. This lowers her expected manipulation costs compared to the separating equilibrium.

The same story applies to divisional managers that report to headquarters, and whose compensation depends on the CEO's perception of their divisional performance.

Capital Allocation Inside the Firm

Capital is allocated inside firms based on the attractiveness of the investment opportunities that exist in various divisions. Divisional managers try to attract more capital to their divisions, so as to increase their compensation and the chances for promotion. Divisional managers provide the headquarters with information regarding the attractiveness of projects in their divisions and in the process may upwardly bias the estimates of their profitability (see, Berkovitch and Israel 2004). Our theory suggests that divisional managers may reveal very good and discard very bad projects, but may pool the quality of the intermediate quality projects. This enables divisional managers to manipulate their reports less, on average, and to reduce the expected penalties.

Raising External Capital

A similar situation may arise when a firm tries to raise external capital. The manager's payoff may very well be tied to firm size, which may cause him to adopt a window-dressing strategy. On the other hand, biasing the reports might result in legal and reputational repercussions. Our equilibrium suggests that while manipulation is inevitable, noise may be endogenously introduced into the reports. While extremely good and extremely bad firms will be revealed, intermediate firms may pool together.

Security Analysts

Sell-side security analysts collect information about the firms they cover, and provide investors with recommendations, earnings forecasts, price targets, etc. Thus, security analysts may be considered as "experts." Morgan and Stocken (2003) use a "cheap-talk" setup to model the reporting environment of security analysts. This results in a step-function equilibrium, which they interpret as analogous to the finite number of recommendation levels used by analysts ("strong buy," "buy," "hold," etc.).

Our model, suggests that a similar result may obtain when analysts face direct cost of misrepresentation.[33] A difference between the two results is that our model suggests perfect revelation for extremely good and extremely bad outcomes, while in the "cheap talk" framework, perfect revelation is impossible on any interval of types.

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[33] Recent regulatory changes suggest that such a direct cost may actually exist. For example, the "Global Analyst Research Settlement" of 2003 imposed severe monetary penalties on the top U.S. brokerage houses due to alleged excess optimism of their sell-side analysts (see Kadan et al. 2005).

Another application may be to earnings forecasts. It is well documented that analysts may have an incentive to bias their forecasts.[34] Analysts also care about the accuracy of their forecast and, hence, biasing the forecast is costly for them.[35] Our model suggests a fundamental difference between "clear" forecasts (revealing, either good or bad) and "not so clear" intermediate forecasts. The model suggests that extremely good or bad forecasts will be fully revealing. Analysts in this case will provide a "bold forecast" that reveals their private information. By contrast, if the information that the analyst possesses falls in an intermediate range, then the analyst will pool his report with reports of other analysts, resulting in something that looks like a "herding" behavior, and in less information being conveyed to the market.

VII. CONCLUSION

Theoretical papers on earnings management have focused on equilibria in which the distribution of reported earnings is smooth (e.g., Fischer and Verrecchia 2000). In this paper, we show that even in a setup where both costs and incentives for manipulation are smooth, there are partially pooling equilibria in which discontinuities in the reports evolve endogenously. We demonstrate the partially pooling equilibria in the context of earnings management, but suggest that other financial reporting environments fit into the basic framework.

In the partially pooling equilibrium, if the earnings are either low or high, then the manager issues a report that reveals his type, as in the standard separating equilibrium. If earnings are in an intermediate range, then the manager always reports the same amount (a pooling report); i.e., the investors can infer only the interval of types to which he belongs, but not his actual type. The comparative statics of the model provide predictions for the determinants of the size and the location of the pooling interval. The analysis indicates that the manager always prefers the partially pooling equilibrium over the separating equilibrium, while the investors' preferences depend on the specification of their utility function. If investors are risk-neutral, then they are indifferent between the two equilibria. If investors are risk-averse, then their preference depends on the size of the risk premium factored into the market price.

We discuss general properties of all possible equilibria in this model, and demonstrate the existence of partially pooling equilibria with multiple pooling intervals. All these equilibria involve a separating range on the left-hand side and right-hand side of the type distribution, and multiple pooling intervals in the middle. A special case is a mix of the standard separating equilibrium on both tails of the distribution combined with step function reporting strategies. This equilibrium is quite similar to those obtained in the "cheap talk" literature (see, Crawford and Sobel 1982), but with a different intuition.

Our partially pooling equilibrium can also be applied to other environments related to financial reporting. We suggest that internal financial reporting within a firm, internal capital allocation processes, the process of raising outside equity, as well as analysts' forecasts and recommendations, can all fit into the basic framework of this model with only slight modifications. In all these cases, an informed expert issues a report to a less informed decision maker. Extremely good and extremely bad recommendations are revealed, while intermediate outcomes are pooled together in a noisy report.

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[34] Dugar and Nathan (1995), Lin and McNichols (1998), Michaely and Womack (1999), Lim (2001), and many others provide evidence about biases of security analysts.

[35] See, for example, Mikhail et al. (1999).

Future research may use our results for both theoretical and empirical purposes. First, equilibria similar to the ones we identified in this paper may be used in future theoretical research to model a variety of phenomena related to costly signaling. Second, our paper yields several empirical predictions regarding the location (relative to the mean), size, and probability of the Burgstahler and Dichev (1997) kinks in the distribution of reported earnings. We expect future empirical research to test these predictions.

APPENDIX
Proofs

Proof of Proposition 1

Let $\rho_s(\cdot)$ be a perfectly separating, continuously differentiable reporting strategy. Since $\rho_s(\cdot)$ is perfectly separating it can be inverted; thus, let $\varphi_s = \rho_s^{-1}$. It follows that the only pricing function consistent with $\rho_s(\cdot)$ is given by $P_s(\cdot) = \varphi_s(\cdot)$. The payoff of the manager given earnings of x and a report of x^R is given by:

$$U^M(x,x^R) = \alpha\varphi_s(x^R) - \beta(x^R - x)^2. \tag{11}$$

The first order condition of (11) with respect to x^R is:

$$\frac{d}{dx^R}\varphi_s(x^R) - \frac{2\beta}{\alpha}x^R + \frac{2\beta}{\alpha}x = 0.$$

Since in equilibrium $x = \varphi_s(x^R)$, we obtain the following linear, first-order differential equation for an equilibrium:

$$\frac{d}{dx^R}\varphi_s(x^R) = -\frac{2\beta}{\alpha}\varphi_s(x^R) + \frac{2\beta}{\alpha}x^R.$$

All potential solutions of this equation are given by:

$$\varphi_s(x^R) = x^R - \frac{\alpha}{2\beta} + Ke^{\frac{2x^R\beta}{\alpha}},$$

where K is a constant. We claim that $K = 0$. Indeed, suppose on the contrary that $K > 0$, then a simple calculation shows that $\varphi_s(x^R)$ is strictly convex and has a unique minimum at $x^R = -\frac{\alpha}{2\beta}\ln\frac{\alpha}{2k\beta}$. This implies that $\varphi_s(x^R)$ is bounded from below, contrary to the fact that x can take any value in R (it is drawn from a normal distribution that is supported on the entire real line). Similarly, we can rule out the case $K < 0$. Therefore, we obtain $\varphi_s(x^R) = x^R - \frac{\alpha}{2\beta}$, $P_s(x^R) = x^R - \frac{\alpha}{2\beta}$ and $\rho_s(x) = x + \frac{\alpha}{2\beta}$, as required. ■

Proof of Proposition 2

The proof consists of two parts. In Part 1 we show that if there exists an interval $[a,b]$ such that the necessary Conditions (6) and (7) are satisfied, then the reporting strategy and pricing function specified in the statement of the proposition applied to this interval are indeed an equilibrium. Part 2 shows that such an interval actually exists, and furthermore it is unique.

Part 1

Assume that there exists a unique interval $[a,b]$ such that the necessary Conditions (6) and (7) are satisfied. In particular, this interval makes the types $x = a$ and $x = b$ indifferent between reporting b and reporting $x + \frac{\alpha}{2\beta}$. We claim that $\rho_p^*(\cdot)$ applied to this interval is an equilibrium strategy.

The pricing function on the equilibrium path satisfies Bayes rule, given the manager's strategy by construction. Since $\rho_s^*(\cdot)$ is an equilibrium, and $\rho_p^*(\cdot)$ differs from $\rho_s^*(\cdot)$ only on $[a,b]$, we only have to rule out deviations to and from the pooling interval $[a,b]$.

Consider first a type $\hat{x} \in (a,b)$. The conjectured equilibrium strategy $\rho_p^*(\cdot)$ specifies that he should report b. Since the out-of-equilibrium pricing is identical to the pricing given $\rho_s^*(\cdot)$, Proposition 1 implies that his best possible deviation is to report $\rho_s^*(\hat{x}) = \hat{x} + \frac{\alpha}{2\beta}$. However, using the facts that $d(a,b) = a + \frac{3\alpha}{4\beta}$, and $b = a + \frac{\alpha}{\beta}$ we obtain:

$$\begin{aligned} U^M(\hat{x},\rho_p^*(\hat{x})) - U^M(\hat{x},\rho_s^*(\hat{x})) &= U^M(\hat{x},b) - U^M\left(\hat{x},\hat{x} + \frac{\alpha}{2\beta}\right) \\ &= \alpha d(a,b) - \beta(b - \hat{x})^2 - \left[\alpha\hat{x} - \beta\left(\frac{\alpha}{2\beta}\right)^2\right] \\ &= \beta(\hat{x} - a)(b - \hat{x}) > 0, \end{aligned} \tag{12}$$

where the inequality follows since $\hat{x} \in (a,b)$. Therefore, type $\hat{x}$ is better-off reporting b as required.

Consider now a type $\hat{x} \notin [a,b]$. According to $\rho_p^*(\cdot)$ he should report $\hat{x} + \frac{\alpha}{2\beta}$. Since the out-of-equilibrium pricing is identical to the pricing function given $\rho_s^*(\cdot)$, Proposition 1 implies that this type would not deviate to any report in $\left[a + \frac{\alpha}{2\beta}, b\right) \cup \left(b, b + \frac{\alpha}{2\beta}\right]$. Thus, his only potential beneficial deviation is to report b. However, using the facts that $d(a,b) = a + \frac{3\alpha}{4\beta}$, and $b = a + \frac{\alpha}{\beta}$ we obtain:

$$\begin{aligned} U^M\left(\hat{x},\hat{x} + \frac{\alpha}{2\beta}\right) - U^M(\hat{x},b) &= \alpha\hat{x} - \beta\left(\frac{\alpha}{2\beta}\right)^2 - \alpha d(a,b) + \beta(\hat{x} - b)^2 \\ &= \alpha\hat{x} - \beta\left(\frac{\alpha}{2\beta}\right)^2 - \alpha\left(a + \frac{3\alpha}{4\beta}\right) + \beta\left(\hat{x} - a - \frac{\alpha}{\beta}\right)^2 \\ &= \beta(\hat{x} - a)(\hat{x} - b) > 0, \end{aligned}$$

where the inequality follows since $\hat{x} \notin [a,b]$. Therefore, no deviation is beneficial. ■

Part 2

It is sufficient to show that there exists a unique $a \in R$, such that $E\left(\tilde{x} | \tilde{x} \in \left[a, a + \frac{\alpha}{\beta}\right]\right) = a + \frac{3\alpha}{4\beta}$. As $b = a + \frac{\alpha}{\beta}$ we shall denote the conditional expectation by $d(a) = E\left(\tilde{x} | \tilde{x} \in \left[a, a + \frac{\alpha}{\beta}\right]\right)$ instead of $d(a,b)$. Thus, we will show that there exists a unique $a \in R$, such that $d(a) = a + \frac{3\alpha}{4\beta}$. The conditional expectation $d(a)$ is the expectation of a truncated normal random variable over the interval $\left[a, a + \frac{\alpha}{\beta}\right]$. It is well known (see Johnson et al. 1994) that $d(a)$ may be expressed using the following formula:

$$d(a) = x_0 - \sigma^2 \frac{f\left(a + \frac{\alpha}{\beta}\right) - f(a)}{F\left(a + \frac{\alpha}{\beta}\right) - F(a)} \quad a \in R. \tag{13}$$

Also, notice that the first derivative of the normal density satisfies:

$$f'(x) = -\frac{x - x_0}{\sigma^2} f(x). \tag{14}$$

The following two lemmas are needed to establish the existence of the required a.

Lemma 2: The following holds for any $s > 0$:

$$\lim_{x \to \infty} \frac{f(x + s)}{f(x)} = 0$$

$$\lim_{x \to -\infty} \frac{f(x + s)}{f(x)} = \infty.$$

Proof: For any $s > 0$, and $x \in R$ we have:

$$\frac{f(x + s)}{f(x)} = \frac{\frac{1}{\sqrt{2\pi\sigma^2}} \exp - \frac{(x + s - x_0)^2}{2\sigma^2}}{\frac{1}{\sqrt{2\pi\sigma^2}} \exp - \frac{(x - x_0)^2}{2\sigma^2}} = \exp - \frac{s(2x - 2x_0 + s)}{2\sigma^2}.$$

The result follows by taking the appropriate limits. ■

Lemma 3: The following holds:

$$\lim_{a \to \infty} [d(a) - a] = 0$$

$$\lim_{a \to -\infty} [d(a) - a] = \frac{\alpha}{\beta}.$$

Proof: By (13), and using L'Hopital's law and the properties of the normal distribution:

$$\lim_{a\to\infty} d(a) = x_0 - \sigma^2 \lim_{a\to\infty} \frac{f\left(a+\frac{\alpha}{\beta}\right) - f(a)}{F\left(a+\frac{\alpha}{\beta}\right) - F(a)} = x_0 - \sigma^2 \lim_{a\to\infty} \frac{f'\left(a+\frac{\alpha}{\beta}\right) - f'(a)}{f\left(a+\frac{\alpha}{\beta}\right) - f(a)}$$

$$= x_0 + \sigma^2 \lim_{a\to\infty} \frac{\frac{a+\frac{\alpha}{\beta}-x_0}{\sigma^2} f\left(a+\frac{\alpha}{\beta}\right) - \frac{a-x_0}{\sigma^2} f(a)}{f\left(a+\frac{\alpha}{\beta}\right) - f(a)}$$

$$= x_0 + \lim_{a\to\infty} \frac{\left(a+\frac{\alpha}{\beta}-x_0\right) f\left(a+\frac{\alpha}{\beta}\right) - (a-x_0) f(a)}{f\left(a+\frac{\alpha}{\beta}\right) - f(a)}$$

$$= \lim_{a\to\infty} \left[a + \frac{\alpha}{\beta} \frac{f\left(a+\frac{\alpha}{\beta}\right)}{f\left(a+\frac{\alpha}{\beta}\right) - f(a)} \right]$$

Now, by plugging $s = \frac{\alpha}{\beta}$ in Lemma 2 it follows that:

$$\lim_{a\to\infty} [d(a) - a] = \frac{\alpha}{\beta} \lim_{a\to\infty} \frac{f\left(a+\frac{\alpha}{\beta}\right)}{f\left(a+\frac{\alpha}{\beta}\right) - f(a)} = \frac{\alpha}{\beta} \lim_{a\to\infty} \frac{1}{1 - \frac{f(a)}{f\left(a+\frac{\alpha}{\beta}\right)}} = 0, \text{ as required.}$$

As for the second part, repeating the previous analysis we obtain:

$$\lim_{a\to-\infty} d(a) = \lim_{a\to-\infty} \left[a + \frac{\alpha}{\beta} \frac{f\left(a+\frac{\alpha}{\beta}\right)}{f\left(a+\frac{\alpha}{\beta}\right) - f(a)} \right].$$

Using Lemma 2 it follows that:

$$\lim_{a \to -\infty} [d(a) - a] = \frac{\alpha}{\beta} \lim_{a \to -\infty} \frac{f\left(a + \frac{\alpha}{\beta}\right)}{f\left(a + \frac{\alpha}{\beta}\right) - f(a)} = \frac{\alpha}{\beta} \lim_{a \to -\infty} \frac{1}{1 - \frac{f(a)}{f\left(a + \frac{\alpha}{\beta}\right)}} = \frac{\alpha}{\beta},$$

as required. ■

It is now easy to prove the existence of a required a. Indeed, define $H(a) \equiv d(a) - a - \frac{3\alpha}{4\beta}$. From Lemma 3 it follows that $\lim_{a \to -\infty} H(a) = \frac{\alpha}{4\beta} > 0$, and $\lim_{a \to \infty} H(a) = -\frac{3\alpha}{4\beta} < 0$. Thus, from the continuity of $H(a)$ there exists an $a \in R$ such that $H(a) = 0$.

The next step is to prove the uniqueness of the chosen a. This proof is technical and lengthy and therefore we do not report it here. It is available upon request. ■

Proof of Corollary 3

For brevity we assume $x_0 = 0$.[36] Since we are interested in the impact of σ, we view a and d as functions of σ. Define:

$$H(a,\sigma) \equiv d(a,\sigma) - a - \frac{3\alpha}{4\beta}.$$

The relation between a and σ is given by the implicit equation $H(a,\sigma) = 0$. In the proof of Proposition 2 we have shown that $\frac{\partial H(a,\sigma)}{\partial a} < 0$ for all $a,\sigma \in R$. By the implicit function theorem we have:

$$\frac{\partial a(\sigma)}{\partial \sigma} = -\frac{\frac{\partial H(a,\sigma)}{\partial \sigma}}{\frac{\partial H(a,\sigma)}{\partial a}}.$$

Thus, to show that $\frac{\partial a(\sigma)}{\partial \sigma} < 0$ it is sufficient to show that $\frac{\partial H(a,\sigma)}{\partial \sigma} < 0$. We have:

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[36] A different choice of x_0 would shift $a(\sigma)$ by a constant, and thus would have no effect on $\frac{\partial a}{\partial \sigma}$.

$$\frac{\partial H(a,\sigma)}{\partial\sigma}=\frac{\partial d(a,\sigma)}{\partial\sigma}=\frac{\partial}{\partial\sigma}\frac{\int_a^{a+\frac{\alpha}{\beta}}xf(x)dx}{F\left(a+\frac{\alpha}{\beta}\right)-F(a)}=\frac{\partial}{\partial\sigma}\frac{\frac{1}{\sqrt{2\pi\sigma^2}}\int_a^{a+\frac{\alpha}{\beta}}xe^{-\frac{x^2}{2\sigma^2}}dx}{\frac{1}{\sqrt{2\pi\sigma^2}}\int_a^{a+\frac{\alpha}{\beta}}e^{-\frac{x^2}{2\sigma^2}}dx}$$

$$=\frac{\int_a^{a+\frac{\alpha}{\beta}}\frac{x^3}{\sigma^3}e^{-\frac{x^2}{2\sigma^2}}dx\cdot\int_a^{a+\frac{\alpha}{\beta}}e^{-\frac{x^2}{2\sigma^2}}dx-\int_a^{a+\frac{\alpha}{\beta}}\frac{x^2}{\sigma^3}e^{-\frac{x^2}{2\sigma^2}}dx\cdot\int_a^{a+\frac{\alpha}{\beta}}xe^{-\frac{x^2}{2\sigma^2}}dx}{\left[\int_a^{a+\frac{\alpha}{\beta}}e^{-\frac{x^2}{2\sigma^2}}dx\right]^2}$$

$$=\frac{1}{\sigma^3}\left[\frac{\int_a^{a+\frac{\alpha}{\beta}}x^3\,f(x)dx}{\int_a^{a+\frac{\alpha}{\beta}}f(x)d\pi}-\frac{\int_a^{a+\frac{\alpha}{\beta}}x^2\,f(x)dx}{\int_a^{a+\frac{\alpha}{\beta}}f(x)dx}\cdot\frac{\int_a^{a+\frac{\alpha}{\beta}}xf(x)dx}{\int_a^{a+\frac{\alpha}{\beta}}f(x)dx}\right]$$

$$=\frac{1}{\sigma^3}\left[E\left(\tilde{x}^3|a\leq\tilde{x}\leq a+\frac{\alpha}{\beta}\right)-E\left(\tilde{x}^2|a\leq\tilde{x}\leq a+\frac{\alpha}{\beta}\right)\right.$$
$$\left.\cdot E\left(\tilde{x}|a\leq\tilde{x}\leq a+\frac{\alpha}{\beta}\right)\right]$$
$$=\frac{1}{\sigma^3}Cov\left(\tilde{x}^2,\tilde{x}|a\leq\tilde{x}\leq a+\frac{\alpha}{\beta}\right).$$

Thus, the sign of $\frac{\partial H(a,\sigma)}{\partial\sigma}$ is equal to the sign of $Cov(\tilde{y}, \tilde{y}^2)$, where $\tilde{y}$ is a random variable obtained from a truncation of $\tilde{x}$ between a and $a + \frac{\alpha}{\beta}$. It can be shown that this covariance is strictly negative, as required.[37] The fact that $d(a)$ decreases in σ follows immediately since $d(a) = a + \frac{3\alpha}{4\beta}$. ■

Given that $a(\sigma)$ is decreasing in σ, we know that $a_0 \equiv \lim_{\alpha\to 0} a(\sigma)$ exists. It is easy to see that for any fixed $a < x_0$ and $b > a$ we have:

$$\lim_{\sigma\to 0} E(\tilde{x}|\tilde{x}\in[a,b])=\begin{cases}b & x_0\notin[a,b]\\ x_0 & x_0\in[a,b]\end{cases}. \tag{15}$$

We claim first that there exists an $a_0 > 0$ such that $a_0 + \frac{\alpha}{\beta} > x_0$. Indeed, suppose on the contrary that $a_0 + \frac{\alpha}{\beta} \leq x_0$. This implies by (15) that $\lim_{\sigma\to 0} E\left(\tilde{x}|\tilde{x}\in\left[a(\sigma),a(\sigma)+\frac{\alpha}{\beta}\right]\right)$

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[37] The proof is technical. It applies to any symmetric and continuous distribution, and not only to the normal distribution. It relies on the fact that the truncation interval $\left[a,a+\frac{\alpha}{\beta}\right]$ is tilted to the left-hand side of the distribution. We omit the proof here for brevity, but it is available upon request.

$= a_0 + \frac{\alpha}{\beta}$, contradicting the fact that for all σ, $E\left(\tilde{x} \mid \tilde{x} \in \left[a(\sigma), a(\sigma) + \frac{\alpha}{\beta}\right]\right) = a(\sigma) + \frac{3\alpha}{4\beta}$. Now, for all $\varepsilon > 0$ sufficiently small we have: $a_0 - \varepsilon + \frac{\alpha}{\beta} > x_0$. Thus, by (15) we have:

$$\lim_{\sigma \to 0} E\left(\tilde{x} \mid \tilde{x} \in \left[a_0 - \varepsilon, a_0 - \varepsilon + \frac{\alpha}{\beta}\right]\right) = x_0.$$

From the continuity of the conditional expectation and since ε is arbitrary we conclude that $\lim_{\sigma \to 0} E\left(\tilde{x} \mid \tilde{x} \in \left[a(\sigma), a(\sigma) + \frac{\alpha}{\beta}\right]\right) = x_0$. And, hence $\lim_{\sigma \to 0} a(\sigma) = x_0 - \frac{3\alpha}{4\beta}$, as required.

As for the case of $\sigma \to \infty$, for all fixed a and b we have: $E(\tilde{x} \mid \tilde{x} \in [a,b]) \to \frac{a+b}{2}$. Indeed, by applying L'Hopital's law we obtain:

$$\lim_{\sigma \to \infty} \left[x_0 - \sigma^2 \frac{f(b) - f(a)}{F(b) - F(a)}\right] = x_0 - \lim_{\sigma \to \infty} \sigma^2 \frac{e^{\frac{-(b-x_0)^2}{2\sigma^2}} - e^{\frac{-(a-x_0)^2}{2\sigma^2}}}{\int_a^b e^{\frac{-(x-x_0)^2}{2\sigma^2}} dx}$$

$$= x_0 - \frac{1}{b-a} \lim_{\sigma \to \infty} \frac{e^{\frac{-(b-x_0)^2}{2\sigma^2}} - e^{\frac{-(a-x_0)^2}{2\sigma^2}}}{\frac{1}{\sigma^2}}$$

$$= x_0 - \frac{1}{b-a} \lim_{\sigma \to \infty} \frac{\frac{(b-x_0)^2}{\sigma^3} e^{\frac{-(b-x_0)^2}{2\sigma^2}} - \frac{(a-x_0)^2}{\sigma^3} e^{\frac{-(a-x_0)^2}{2\sigma^2}}}{-\frac{2}{\sigma^3}}$$

$$= x_0 + \frac{(b-x_0)^2 - (a-x_0)^2}{2(b-a)} = \frac{a+b}{2}$$

This calculation implies that if $a_\infty \equiv \lim_{\sigma \to \infty} a(\sigma)$ were finite, then we would have that $d(a(\sigma)) \to a_\infty + \frac{\alpha}{2\beta}$—a contradiction to the fact $d(a(\sigma)) = a(\sigma) + \frac{3\alpha}{4\beta}$ for all σ. ■

Proof of Proposition 3

The expected earnings management in the separating equilibrium is $\frac{\alpha}{2\beta}$. The expected earnings management in the partially pooling equilibrium is:

$$\int_{-\infty}^{a} \frac{\alpha}{2\beta} f(x)dx + \int_{a}^{b} (b - x)f(x)dx + \int_{b}^{\infty} \frac{\alpha}{2\beta} f(x)dx \tag{16}$$
$$= \frac{\alpha}{2\beta} - \frac{\alpha}{2\beta}(F(b) - F(a)) + \int_{a}^{b} (b - x)f(x)dx$$
$$= \frac{\alpha}{2\beta} - \frac{\alpha}{2\beta}(F(b) - F(a)) + b(F(b) - F(a)) - d(a,b)(F(b) - F(a))$$
$$= \frac{\alpha}{2\beta} - \frac{\alpha}{4\beta}(F(b) - F(a)),$$

where the second equality follows since by definition: $d(a,b) = \frac{\int_a^b xf(x)dx}{F(b) - F(a)}$, and the last equality follows from the fact that $b - d(a,b) = \frac{\alpha}{4\beta}$ (Equations (6) and (7)). This implies that the expected earnings management in the pooling equilibrium is strictly lower than in the separating equilibrium. ■

Proof of Proposition 4

The inequality follows directly from Equation (12). Indeed, this inequality states that for every realized earnings in the pooling interval, the payoff of the manager is strictly higher under the partially pooling equilibrium.

Proof of Proposition 6

To prove this proposition note first that for any given pricing function $P(\cdot)$ the Spence-Mirrlees single crossing property holds. That is:

$$\frac{\partial^2 U^M(x,x^R)}{\partial x^R \partial x} = 2\beta > 0 \text{ for all } x \text{ and } x^R. \tag{17}$$

Part 1

Let $P(\cdot)$ be any pricing function. For any x denote:

$$BR(x,P) \equiv \arg\max_{x^R} \{\alpha P(x^R) - \beta(x - x^R)^2\}.$$

Then $BR(x)$ is the best response of type x given the pricing function. It may consist of one point (a singleton) or many points. Let $x_2 > x_1$, and let $P(\cdot)$ be any pricing function, we claim that $\inf BR(x_2,P) \geq \sup BR(x_1,P)$.

To see this, let $y_1 = \sup BR(x_1,P)$ and $y_2 = \inf BR(x_2,P)$. Suppose on the contrary that $y_2 < y_1$. Then, there exists an $x_1^R \in BR(x_1)$ and $x_2^R \in BR(x_2)$ with $x_1^R > x_2^R$.

Since $x_1^R \in BR(x_1,P)$ we know that $U^M(x_1,x_1^R) \geq U^M(x_1,x_2^R)$. From (17) we conclude that $U^M(x_2,x_1^R) > U^M(x_2,x_2^R)$. But, this contradicts the fact that $x_2^R \in BR(x_2)$. This establishes that best responses are non-decreasing for any pricing function. But, equilibrium reporting functions are best responses to the equilibrium pricing functions, and so must be non-decreasing.

Part 2

Any non-decreasing function has at most a countable number of discontinuities, and hence the set of types for which the reporting function is discontinuous has probability 0.

Part 3

Consider any equilibrium with a pricing function $P(\cdot)$. Let M be the set of types that use a mixed strategy in this equilibrium. For any $x \in M$, $BR(x,P)$ must consist of more than one point, because this types randomizes between several reports. For any $x \in M$ define: $a_x \equiv \inf BR(x,P)$, $b_x \equiv \sup BR(x,P)$. Since the best response correspondence is non-decreasing (Part 1), the open intervals $\{(a_x,b_x)\}_{x \in M}$ are disjoint and non-empty. Hence, each one of these intervals contains a rational number. This implies that there are at most a countable number of such intervals. But then, this implies that M is at most countable and, hence, has a zero measure.

Part 4

Suppose, for example that ρ is discontinuous at $x = \hat{x}$, but ρ is not constant either to the right or to the left of $\hat{x}$ (that is there is no pooling interval adjacent to $\hat{x}$). Then, ρ must be strictly increasing in the neighborhood of $\hat{x}$, since it is non-decreasing (by Part 1). In particular, it is separating in the neighborhood of $\hat{x}$ and, hence, the stock price is strictly increasing in the report and reflects the true type. But then one can choose $\varepsilon > 0$ small enough such that a type $\hat{x} + \varepsilon$ will benefit from deviating by reporting $\rho(\hat{x} - \varepsilon)$. Indeed, the price decrease is on the order of ε and hence can be made arbitrarily small, but the savings in manipulation costs are strictly positive due to the discontinuity of the reports at $\hat{x}$. ■

Proof of Corollary 5

From Proposition 6 we know that the set of types that randomize has probability 0. Hence, it is impossible that all types randomize. The only way a "babbling" equilibrium can occur is if all types provide the same report $\hat{x}^R$. Let $P(x^R)$ be the equilibrium pricing function. Following a report of $\hat{x}^R$ nothing is learned by the investors, and so $P(\hat{x}^R) = x_0$. Other reports $x^R \neq \hat{x}^R$ are not expected in equilibrium, and so we have out of equilibrium pricing.

Note that for every pricing function $P(\cdot)$ and for every report (assume w.l.o.g. that $x^R > \hat{x}^R$) there exists some type $x' > \hat{x}^R$ such that all types higher than x' prefer deviating to x^R over reporting the alleged equilibrium report $\hat{x}^R$. This holds because the marginal cost of manipulation is not bounded. Thus, no out-of-equilibrium beliefs support this equilibrium. ■

Proof of Lemma 1

Recall that given $\eta \in \left(-\frac{\alpha}{2\beta},\frac{\alpha}{2\beta}\right)$, the conditional mean given the pooling report is $d_\eta = b_\eta - \frac{\alpha}{4\beta} + \frac{\beta}{\alpha}\eta^2$, and the size of the pooling interval is $b_\eta - a_\eta = \frac{\alpha}{\beta} - 2\eta$. Denote:

$$r(\eta) \equiv \frac{d_\eta - a_\eta}{b_\eta - a_\eta}, \tag{18}$$

the relative location of the conditional mean in the pooling interval. Then:

$$r(\eta) = \frac{d_\eta - b_\eta + b_\eta - a_\eta}{b_\eta - a_\eta} = 1 - \frac{\frac{\alpha}{4\beta} - \frac{\beta}{\alpha}\eta^2}{\frac{\alpha}{\beta} - 2\eta}. \tag{19}$$

Plugging $\eta = -\frac{\alpha}{2\beta}$ we obtain $r\left(-\frac{\alpha}{2\beta}\right) = 1$. Hence, as η approaches $-\frac{\alpha}{2\beta}$, the pooling interval moves to the far left side of the distribution and hence $\lim_{\eta \to -\frac{\alpha}{2\beta}} a_\eta = -\infty$. Also, using L'Hopital's rule $\lim_{\eta \to \frac{\alpha}{2\beta}} r(\eta) = \frac{1}{2}$. And so, as η tends to $\frac{\alpha}{2\beta}$ the conditional mean d_η tends to the unconditional mean x_0. But, since $\lim_{\eta \to \frac{\alpha}{2\beta}} (b_\eta - a_\eta) = 0$, we get that $\lim_{\eta \to \frac{\alpha}{2\beta}} a_\eta = x_0$. From continuity we get that a_η takes all possible values smaller than x_0.

■

REFERENCES

Abarbanell, J., and R. Lehavy. 2003. Biased forecasts or biased earnings? The role of reported earnings in explaining apparent optimism and inefficiency in analysts' earnings forecasts. *Journal of Accounting and Economics* 36: 105–146.

Baiman, S., H. Evans, and J. Noel. 1987. Optimal contracts with a utility maximizing auditor. *Journal of Accounting Research* 25: 217–244.

Barnea, A., J. Ronen, and S. Sadan. 1975. The implementation of accounting objectives—An application to extraordinary items. *The Accounting Review* 50: 58–68.

Beaver, W., M. McNichols, and K. Nelson. 2004. An alternative interpretation of the discontinuity in earnings distributions. Working paper, Stanford University.

Berkovitch, E., and R. Israel. 2004. Why the NPV criterion does not maximize NPV. *Review of Financial Studies* 17: 239–255.

Bergstresser, D., and T. Philippon. 2006. CEO incentives and earnings management. *Journal of Financial Economics* 80: 511–529.

Burgstahler, D., and I. Dichev. 1997. Earnings management to avoid earnings decreases and losses. *Journal of Accounting and Economics* 24: 99–126.

Burns, N., and S. Kedia. 2006. Do executive stock options generate incentives for earnings management? Evidence from accounting restatements. *Journal of Financial Economics* 79: 35–67.

Cheng, Q., and T. Warfield. 2005. Equity incentives an earnings management. *The Accounting Review* 80: 441–476.

Crawford, V., and J. Sobel. 1982. Strategic information transmission. *Econometrica* 50: 1431–1451.

Dechow, P. M., S. A. Richardson, and A. I. Tuna. 2003. Why are earnings kinky? An examination of the earnings management explanation. *Review of Accounting Studies* 8: 355–384.

Degeorge, F., J. Patel, and R. Zeckhauser. 1999. Earnings management to exceed thresholds. *Journal of Business* 72: 1–33.

Demski, J., and R. A. Dye. 1999. Risk, return and moral hazard. *Journal of Accounting Research* 37: 27–56.

Dugar, A., and S. Nathan. 1995. The effect of investment banking relationships on financial analysts' earnings forecasts and investment recommendations. *Contemporary Accounting Research* 12: 131–160.

Durtschi, C., and P. Easton. 2005. Earnings management? The shapes of the frequency distributions of earnings metrics are not evidence ipso facto. *Journal of Accounting Research* 43: 557–592.

Dye, R. 1988. Earnings management in an overlapping generations model. *Journal of Accounting Research* 26: 195–235.

———. 2002. Classifications manipulation and Nash accounting standards. *Journal of Accounting Research* 40: 1125–1162.

Evans, J., and S. Sridhar. 1996. Multiple control systems, accrual accounting and earnings management. *Journal of Accounting Research* 34: 45–66.

Ewert, R., and A. Wagenhofer. 2005. Economic effects of tightening accounting standards to restrict earnings management. *The Accounting Review* 80: 1101.

Fischer, P. E., and R. E. Verrecchia. 2000. Reporting bias. *The Accounting Review* 75: 229–245.

Fudenberg, D., and J. Tirole. 1991. Perfect Bayesian and sequential equilibrium. *Journal of Economic Theory* 53: 236–260.

———, and ———. 1995. A theory of income and dividend smoothing based on incumbency rents. *Journal of Political Economy* 103: 75–93.

Gonedes, N. J. 1972. Income smoothing behavior under selected stochastic processes. *Journal of Business* 45: 570–584.

Goel, A., and A. V. Thakor. 2003. Why do firms smooth earnings? *Journal of Business* 76: 151–192.

Guttman, I., Kadan O., and E. Kandel. 2006. A model of professional advice. Working paper, Stanford University.

Healy, P. M. 1985. The effect of bonus schemes on accounting decisions. *Journal of Accounting and Economics* 7: 85–107.

———, and J. Wahlen. 1999. A review of the earnings management literature and its implications for standard setting. *Accounting Horizons* 13: 365–383.

Johnson, N. L., S. Kotz, and N. Balakrishnan. 1994. *Continuous Univariate Distributions*. 2nd edition. New York, NY: Wiley & Sons.

Kadan, O., Madureira L., Wang R., and T. Zach. 2005. Conflicts of interest and stock recommendations—The effect of the global settlement and related regulations. Working paper, Washington University in St. Louis.

———, and J. Yang. 2005. Executive stock options and earnings management, a theoretical and empirical analysis. Working paper, Washington University in St. Louis.

Kreps, D. M., and R. Wilson. 1982. Sequential equilibria. *Econometrica* 50: 863–894.

Lambert, R. A. 1984. Income smoothing as rational equilibrium behavior. *The Accounting Review* 59: 604–18.

———. 2001. Contracting theory and accounting. *Journal of Accounting and Economics* 32: 3–87.

Lim, T. 2001. Rationality and analysts forecast bias. *Journal of Finance* 56: 369–385.

Lin, H., and M. McNichols. 1998. Underwriting relationships, analysts' earnings forecasts and investment recommendations. *Journal of Accounting and Economics* 25: 101–127.

Michaely, R., and K. Womack. 1999. Conflict of interest and credibility of underwriter analyst recommendations. *Review of Financial Studies* 12: 653–686.

Mikhail, M., B. Walther, and R. Willis. 1999. Does forecast accuracy matter to security analysts? *The Accounting Review* 74: 185–200.

Morgan, J., and P. Stocken. 2003. An analysis of stock recommendations. *RAND Journal of Economics* 34: 183–203.

Newman, P., and R. Sansing. 1993. Disclosure policies with multiple users. *Journal of Accounting Research* 31: 92–112.

Rangan, S. 1998. Earnings management and the performance of seasoned equity offerings. *Journal of Financial Economics* 50: 101–122.

Riley, J. G. 1979. Informational equilibrium. *Econometrica* 47: 331–359.

Ronen, J., and S. Sadan. 1981. *Smoothing Income Numbers: Objectives, Means, and Implications*. Reading, MA: Edison-Wesley Publishing Company.

Sankar, M., and K. R. Subramanyam. 2001. Reporting discretion and private information communication through earnings. *Journal of Accounting Research* 39: 365–386.

Shivakumar, L. 2000. Do firms mislead investors by overstating earnings before seasoned equity offerings? *Journal of Accounting and Economics* 29: 339–371.

Stein, J. C. 1989. Efficient capital markets, inefficient firms: A model of myopic corporate behavior. *Quarterly Journal of Economics* 104: 655–669.

Teoh, S. H., I. Welch, and T. J. Wong. 1998. Earnings management and the long-run market performance of initial public offerings. *Journal of Finance* 53: 1935–1974.

Trueman, B., and S. Titman. 1988. An explanation for accounting income smoothing. *Journal of Accounting Research* 26 (Supplement): 127–139.

Verrecchia, R. 1986. Managerial discretion in the choice among financial reporting alternatives. *Journal of Accounting and Economics* 8: 175–196.

———. 2001. Essays on disclosure. *Journal of Accounting and Economics* 32: 97–180.

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Organizational Slack in Decentralized Firms: The Role of Business Unit Controllers

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ABSTRACT: We study the determinants of organizational slack in large decentralized firms and focus in particular on how management accounting systems (represented by business unit controllers) affect slack. We rely on an adverse selection model to derive several predictions and to motivate our tests. Consistent with this framework, we find that organizational slack (measured by achievability of business unit managers' performance targets) is higher in settings where business unit controllers focus relatively more on providing decision-making information to business unit managers than on providing information for corporate control. We also find that organizational slack is persistent over time and positively associated with business unit growth, our proxy for the extent of information asymmetry between corporate headquarters and local business unit management.

Keywords: *organizational slack; information asymmetry; business unit controller.*

Data Availability: *Data used in this study cannot be made public due to confidentiality agreements with participating firms.*

I. INTRODUCTION

A fundamental theme in management accounting is that firms use accounting systems to serve two broad objectives: facilitate decision making and mitigate control problems.[1] A ubiquitous control problem in most decentralized organizations is "slack," often defined as inefficient use of resources or underperformance by informationally advantaged individuals such as local business unit managers (Antle and Eppen 1985; Young 1985; Merchant 1985; Chow et al. 1988; Waller 1988). We study the determinants of

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[1] Demski and Feltham (1976), Baiman and Demski (1980), Christensen (1981), Baiman and Evans (1983), Penno (1984), Baiman and Sivaramakrishnan (1991), and Antle and Fellingham (1995) are examples of some early studies that highlight distinct control and decision-facilitating roles for accounting information.

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organizational slack in a sample of 104 business units (BUs) of large decentralized firms and focus in particular on how management accounting systems affect organizational slack.

Our framework is a standard adverse selection model in the spirit of Baron and Myerson (1982) and Laffont and Tirole (1986). In particular, we assume a BU manager in a decentralized firm is privately informed about a productivity parameter relevant to BU's operations. This gives rise to a control problem that manifests in the form of organizational slack. Within this framework, we examine the role of a management accounting system that generates two types of information: information that alleviates the control problem by reducing a BU manager's information advantage vis-à-vis corporate headquarters, and decision-making information that improves everyone's understanding of BU operations.[2] We find that, although the latter type of information can be beneficial, such information can also increase organizational slack.

We also extend this framework to a two-period setting, following the approach in Baron and Besanko (1984). In this setting, we find that organizational slack is serially correlated if a BU manager's private information about a productivity parameter in an earlier period is informative about a similar productivity parameter in a subsequent period.

Our adverse selection framework suggests the following empirical hypotheses. First, organizational slack is higher in settings where the information asymmetry between corporate headquarters and BU management is high. Second, organizational slack is higher in BUs where management accounting systems focus relatively more on providing decision-making information than on providing information for corporate control. Third, organizational slack is persistent (positively serially correlated) over time.

Since both organizational slack and its determinants—the extent of information asymmetry in an organization, and informational properties of management accounting systems—are not directly observable, we employ the following empirical proxies. For organizational slack we use achievability of annual performance targets as perceived by BU managers. This is based on the intuition that easily achievable targets are typically associated with BU underperformance and informational rents for BU managers. As indicators of high information asymmetry between corporate headquarters and BU management, we use measures of BU growth and BU environmental uncertainty. This is based on the intuition that a fast-changing BU environment or a BU strategy that emphasizes growth and market share (as opposed to more immediate profits and returns) is likely associated with greater local expertise and knowledge.

The proxy for a management accounting system in our model is the institution of the BU controller. This proxy is motivated by evidence suggesting that BU controllers' duties and responsibilities effectively mirror the dual objectives of management accounting systems. For example, a number of studies suggest that BU controllers typically report to a corporate chief financial officer to facilitate overall planning and control, and at the same time they report to local BU managers to support decision making within a BU (e.g., Sathe 1978; San Miguel and Govindarajan 1984; Granlund and Lukka 1998a). Similarly, Jablonsky and Keating (1998) describe a BU controller both as a "business advocate" (thinking in terms of service and involvement, encouraging wide use of financial information, integrating business operations, etc.) and a "corporate policeman" (thinking in terms of oversight and surveillance, enforcing compliance with policies and procedures, and administering rules and regulations).[3]

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[2] We do not study the optimal choice of management accounting systems. Instead, we assume the firm is exogenously endowed with a management accounting system and examine how the two types of information it provides affect organizational slack.

Our interest in the role of the BU controller is also motivated by an important early study on organizational slack by Onsi (1973) who argues that an essential prerequisite for the existence of organizational slack is that: "[T]he divisional controller has become a viable member of the divisional team by doing his part to facilitate the creation, distribution and management of divisional slack, [which] represents a shift towards divisional interest in the controller's dual role." In this spirit, we use the term "controller BU focus" to describe the extent to which BU controllers focus relatively more on their local duties as opposed to their corporate control responsibilities.[4]

Overall, the results are consistent with our three hypotheses and robust to a number of sensitivity checks. Our findings contribute to the prior literature in two ways. First, we provide evidence on the role of an important management accounting institution (the BU controller) in decentralized firms where organizational slack is likely to be an important control problem. While prior literature has provided descriptive evidence of BU controllers' duties and responsibilities, ours is the first systematic study of the role of BU controllers in alleviating or exacerbating organizational slack. Second, we contribute to the literature on how firms set performance standards and targets to motivate managers' to increase firm value (e.g., Milgrom and Roberts 1992; Laffont and Tirole 1993; Baiman and Rajan 1998). In particular, our finding that organizational slack, measured by BU managers' perceived achievability of their performance targets, is positively serially correlated suggests that firms do not fully adjust performance targets for BU managers' past performance—otherwise, the achievability of performance targets in the future would be independent of past performance relative to target.[5] This evidence supports theoretical arguments that commitments to use past information less aggressively in setting future standards improves managerial incentives to increase firm value.

The rest of the paper is organized as follows. In Section II, we present an adverse selection model that provides a framework for our hypotheses. In Section III, we describe our research methods including our empirical variables of interest. In Section IV, we present our results. Finally in Section V, we offer a brief summary and conclusion.

II. THEORY AND HYPOTHESES

Baseline Model

To guide our study of organizational slack and its link to management accounting systems in decentralized firms, we rely on a standard adverse selection model where a principal (hereafter, corporate headquarters) delegates responsibility for operating a business unit to a privately informed agent (hereafter, BU manager). We adopt the convention that operating costs of the BU are initially borne by the BU manager who is then reimbursed for those costs along with informational rents that may accrue. In particular, we assume corporate headquarter's utility is:

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[3] In a similar vein, in a study conducted by Institute of Management Accountants (1999), BU controllers describe their roles as follows: "I call it wearing two hats. But you have to wear the hat of the management team you are on to help them achieve the goals. You also have to wear the corporate finance hat to recognize that you have a duty, an overall duty to the corporate finance function to do things right and do them correctly, to stand up for what has to be done, and to say no when no is the appropriate thing to say."

[4] There is no consensus regarding terminology in the prior literature. Controller BU focus has the same meaning as "decentralization in the controller's department" (Simon et al. 1954), "a shift towards divisional interest in the dual role of controllers" (Onsi 1973), "service-aid role of the controller" (Hopper 1980), low "controller independence" (San Miguel and Govindarajan 1984), and "business advocate role of the controller" (Jablonsky and Keating 1998). Also, we view "controller BU focus" as a characteristic of management accounting systems. Finally, we acknowledge that BU controllers may behave strategically in response to various organizational incentives, but view such a consideration beyond the scope of this paper.

[5] Indjejikian and Nanda (2002) reach a similar conclusion based on a seven-year comparison of actual bonuses with target bonuses for a sample of senior executives of U.S. firms.

$$v = \beta x - w \tag{1}$$

where x is BU's observable output, w is the BU manager's compensation, and $\beta \in [\underline{\beta}, \bar{\beta}]$ (with $\underline{\beta} \geq 0$) is a random variable representing a common source of uncertainty about the BU's marginal revenue product. In turn, we let the BU manager's utility be:

$$u = w - C(x,\theta) \tag{2}$$

where $C(x,\theta) = \frac{1}{2}x^2\theta$ is the cost of producing x, and $\theta \in [\underline{\theta},\bar{\theta}]$ (with $\underline{\theta} \geq 0$) represents a manager-specific productivity parameter. We assume θ is privately known by the BU manager prior to employment and, thus, is the source of the agency conflict in our model. We assume β and θ are independent.

We assume that, immediately prior to contracting, both corporate headquarters and the BU manager observe a set of signals **s**. We assume **s** represents a management accounting system that is designed to serve two objectives. First, it facilitates BU decision making by providing information about β to both corporate headquarters and the BU manager. Second, it alleviates the control problem by providing corporate headquarters with information about θ (i.e., reduces the BU manager's information advantage about θ vis-à-vis corporate headquarters). We characterize the management accounting system in more detail in the next subsection.

The timeline is as follows (see also Appendix A). Prior to contracting, the BU manager knows θ, whereas corporate headquarter's beliefs about θ are represented by the density $f(\theta|\mathbf{s})$ and *CDF* $F(\theta|\mathbf{s})$. In turn, both the BU manager's and corporate headquarter's beliefs about β are represented by the density $g(\beta|\mathbf{s})$ and CDF $G(\beta|\mathbf{s})$. Subsequently, corporate headquarters offers a contract, (w,x). Given the contract, his private knowledge of θ, and beliefs about β, the BU manager produces output x. Given production costs equal $\frac{1}{2}x^2\theta$, high θ implies less efficient production, and the single crossing property, i.e., $C_{x\theta}(x,\theta) = x > 0$, holds. Hence, the Revelation Principle applies and corporate headquarters effectively asks the BU manager to report his (privately observed) productivity; the BU manager reports θ (based on his private knowledge of θ and beliefs about β), which determines the output x and his wage w.

We state corporate headquarter's unconstrained optimization problem at the time of contracting (conditional on **s**) as follows.[6]

$$\underset{x}{\text{Max}} \int_{\theta} [E(\beta|\mathbf{s})x(\theta,\mathbf{s}) - \tfrac{1}{2}x(\theta,\mathbf{s})^2(\theta + H(\theta|\mathbf{s}))]f(\theta|\mathbf{s})d\theta \tag{3}$$

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[6] We omit detailed derivations since they follow from standard adverse selection models of Baron and Myerson (1982) and others. In brief, incentive compatibility requires $C_{x\theta} > 0$ and $dx^*/d\theta < 0$. Sufficient conditions to ensure $dx^*/d\theta < 0$ are jointly $C_{x\theta} > 0$, $C_{x\theta\theta} \geq 0$, the concavity (w.r.t. x) of the virtual surplus in Equation (3), and the monotone hazard rate condition (i.e., $dH(\theta|\mathbf{s})/d\theta \geq 0$.

where s characterizes the management accounting system, $E(\beta|s) = \int_{\beta} \beta g(\beta|s)d\beta$, and $H(\theta|s) = F(\theta|s)/f(\theta|s)$ is the inverse hazard rate (assumed to be increasing in θ). As in most adverse selection models of this sort, $H(\theta|s)$ reflects the severity of the agency conflict where $H(\theta|s) = 0$ if corporate headquarters perfectly observes the BU managers' type (e.g., Laffont and Tirole 1993; Salanié 1997). Given our assumptions, the second-best production schedule $x^{SB}(\theta,s)$ that solves Equation (3) is:

$$x^{SB} = \frac{E(\beta|s)}{\theta + H(\theta|s)} \tag{4}$$

and the corresponding corporate profit $\pi^{SB}(\theta,s)$ is:

$$\pi^{SB} = \frac{E(\beta|s)^2}{2[\theta + H(\theta|s)]}. \tag{5}$$

In contrast, when θ is observable, the first-best production schedule and corporate profit are higher and given by $x^{FB} = E(\beta|s)/\theta$ and $\pi^{FB} = E(\beta|s)^2/2\theta$, respectively.

We define *organizational slack* as the loss (corporate profits forgone) associated with second-best production relative to first-best production. Hence, we have:

$$SLACK = \frac{E(\beta|s)^2}{2}\left(\frac{1}{\theta} - \frac{1}{\theta + H(\theta|s)}\right). \tag{6}$$

We note that, for a given θ-type manager, greater information asymmetry between corporate headquarters and the BU manager in the sense of high $H(\theta|s)$ implies more organizational slack. To illustrate, consider the following parametric example. Assume there is no management accounting system (i.e., $s = \{\varnothing\}$), and let $f(\theta)$ be uniform over the interval $[\underline{\theta},\bar{\theta}]$. Hence, $H(\theta|s) = H(\theta) = \theta - \underline{\theta}$. For a BU manager of a given type θ, we note that a lower $\underline{\theta}$ implies a higher $H(\theta)$, more underproduction, and hence more organizational slack. Thus, Expression (6) suggests the following hypothesis.

> **H1:** All else equal, organizational slack is higher in environments characterized by greater information asymmetry between corporate headquarters and BU management.

Management Accounting Information

Firms design management accounting systems to serve the dual objectives of decision making and control. In our context, we characterize *decision-making information* as information about β provided to both corporate headquarters and the BU manager. In contrast, we characterize *control information* as information about the BU manager's privately known productivity parameter θ provided to corporate headquarters. That is, while information about β is valuable to corporate headquarters even in the absence of an agency conflict (i.e., even if production decision were made by corporate headquarters), information about θ is valuable only because it reduces the BU manager's information advantage vis-à-vis corporate headquarters.

In particular, let $\mathbf{s} = \{\mathbf{s}_\beta, \mathbf{s}_\theta\}$ represent the management accounting system where the subset $\mathbf{s}_\beta$ is a set of signals informative about β, while $\mathbf{s}_\theta$ (independent of $\mathbf{s}_\beta$) is geared toward the observation of θ. We note that the *ex ante* design of such a management accounting system requires that corporate headquarters maximize the expectation of Equation (5), i.e., $E_s(\pi^{SB})$, subject to the costs of implementing such a system. In what follows, we neither model the choice of an accounting system nor do we focus on the trade-offs involved in such decisions.[7] Instead, we focus on how a given $\mathbf{s}_\beta$ and $\mathbf{s}_\theta$ affect organizational slack. Rewriting Expression (6), we have:

$$SLACK = \frac{E(\beta|\mathbf{s}_\beta)^2}{2}\left(\frac{1}{\theta} - \frac{1}{\theta + H(\theta|\mathbf{s}_\theta)}\right). \tag{6'}$$

Expression (6′) suggests organizational slack is affected by both decision-making information $\mathbf{s}_\beta$ and control information $\mathbf{s}_\theta$. For example, holding $\mathbf{s}_\beta$ fixed, if more control information (via $\mathbf{s}_\theta$) decreases $H(\theta|\mathbf{s}_\theta)$, then it reduces organizational slack. Similarly, holding $\mathbf{s}_\theta$ fixed, if decision-making information increases the expected marginal revenue, $E(\beta|\mathbf{s}_\beta)$, then it also increases organizational slack. This is because, even though corporate profits are higher with decision-making information $\mathbf{s}_\beta$, the profits forgone (i.e., relative to first-best) are higher as well.

Consider the following parametric example. Let $\mathbf{s} = (\mathbf{s}_\beta, \mathbf{s}_\theta)$ be $m + n$-dimensional where $\mathbf{s}_\beta$ represents a sample of m independent observations about β and $\mathbf{s}_\theta$ represents a sample of n independent observations about θ. Also, let $\mathbf{s}_\beta$ be such that $g(\beta|\mathbf{s}_\beta)$ is uniform with $\beta \in [\max(\underline{\beta}, s_1, \ldots, s_m), \overline{\beta}]$.[8] That is, assume that information about β increases the marginal revenue the BU expects to earn. It follows that $E(\beta|\mathbf{s}_\beta) = [\max(\underline{\beta}, s_1, \ldots, s_m) + \overline{\beta}]/2$ which is weakly increasing in m, the number of observations about β. Hence, organizational slack in (6′) is (weakly) increasing in m. In a similar vein, assume that $\mathbf{s}_\theta$ is such that $f(\theta|\mathbf{s}_\theta)$ is uniform with $\theta \in [\max(\underline{\theta}, s_1, \ldots, s_n, \overline{\theta}]$. It follows that $H(\theta|\mathbf{s}_\theta) = \theta - \max(\underline{\theta}, s_1, \ldots, s_n)$ which is weakly decreasing in n, the number of observations about θ. Hence, organizational slack in (6′) is (weakly) decreasing in n.

This example and the preceding discussion suggest that while information about θ mitigates the BU manager's information advantage and, hence, limits organizational slack, information about β may exacerbate organizational slack. This suggests the following hypothesis.

H2: Organizational slack is higher in settings where the managerial accounting system emphasizes decision-making information at the expense of control information (i.e., m is higher and/or n is lower).

A Two-Period Extension

To illustrate the persistence of organizational slack through time, we consider a two-period adverse selection model following the framework of Baron and Besanko (1984). In particular, we let $v_t = \beta_t x_t - w_t$ and $C_t = \frac{1}{2}x_t^2\theta_t$, $t = 1, 2$, and simplify the setting considerably by assuming that (β_1,β_2) are known and moreover that there is no management accounting system (a timeline is provided in Appendix A). In addition, we assume corporate headquarters commits to a two-period contract (w_1,w_2), and we let the BU manager's pre-contract knowledge of θ_1 be informative about θ_2. That is, we assume θ_1 is positively correlated with θ_2, which implies that a BU manager who knows he is "very productive" in the first period (i.e., low θ_1) believes he is likely to be very productive in the second period as well. Thus, if the density $f_2(\theta_2)$ and CDF $F_2(\theta_2)$ represents (common) prior beliefs about θ_2, then a BU manager who knows θ_1 prior to contracting will form beliefs $f_2(\theta_2|\theta_1)$ and $F_2(\theta_2|\theta_1)$ where $\partial F_2(\theta_2|\theta_1)/\partial\theta_1 \leq 0$. Given our assumptions, it is straightforward to show that the second-best production schedule for the two periods, (x_1^{SB},x_2^{SB}), is given by the following:[9]

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[7] That said, one can easily construct parametric examples where $E_s(\pi^{SB})$ is increasing in both decision-making and control information. Hence, to the extent that firms have limited resources, an optimal trade-off between decision-making and control information is called for. In some settings, providing decision-making information may be paramount, while in others generating control information may be a higher priority.

[8] This follows if the joint distribution of $(\beta, \mathbf{s}_\beta)$ is $(m + 1)/(\overline{\beta} - \underline{\beta})^{m+1}$ with $\underline{\beta} \leq \max(\underline{\beta}, s_1, \ldots, s_i, \ldots, s_m) \leq \beta$ and $\underline{\beta} \leq \beta \leq \overline{\beta}$.

$$x_1^{SB} = \frac{E(\beta_1)}{\theta_1 + H(\theta_1)}, \qquad x_2^{SB} = \frac{E(\beta_2)}{\theta_2 + K(\theta_1,\theta_2)H(\theta_1)}, \tag{7}$$

where $K(\theta_1,\theta_2) = -\frac{\partial F_2(\theta_2|\theta_1)/\partial\theta_1}{f_2(\theta_2|\theta_1)} \geq 0$.

From Equation (7) we note that if θ_1 and θ_2 are independent (i.e., $\partial F_2(\theta_2|\theta_1)/\partial\theta_1 = 0$), then $K(\theta_1,\theta_2) = 0$ and second period production is at first-best levels. This follows since at the time of contracting both corporate headquarters and the BU manager are symmetrically informed about the second period. On the other hand, if θ_1 is informative about θ_2, then the adverse selection problem also infects the second period. As in the one-period model, we define organizational slack in each period as the loss (corporate profits forgone) associated with second-best production relative to first-best production. We have:

$$SLACK_1 = \frac{\beta_1^2}{2}\left(\frac{1}{\theta_1} - \frac{1}{\theta_1 + H(\theta_1)}\right), \text{ and } SLACK_2 = \frac{\beta_2^2}{2}\left(\frac{1}{\theta_2} - \frac{1}{\theta_2 + K(\theta_1,\theta_2)H(\theta_1)}\right). \tag{8}$$

We note that high $H(\theta_1)$, due, for example, to greater information asymmetry between corporate headquarters and a θ_1-type BU manager implies more organizational slack in the first period *and* more organizational slack in the second period. This suggests the following hypothesis.[10]

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[9] We omit detailed derivations since they are straightforward adaptations of results in Baron and Besanko (1984). In brief, incentive compatibility requires $\frac{\partial^2 C_t}{\partial x_t \partial\theta_t} > 0$ and $\frac{dx_t^*}{d\theta_t} < 0$ for $t = 1,2$, and $\lambda E\left(\frac{\partial F_2(\theta_2|\theta_1)\ \partial\theta_1}{f_2(\theta_2|\theta_1)} \frac{\partial^2 C_2}{\partial x_2 \partial\theta_2} \frac{dx_2^*}{d\theta_1}\right) \geq 0$ where the expectation is with respect to θ_2 and λ is the time discount factor. The first two conditions are the standard single-crossing property and the monotonicity conditions applied to each period t. $\frac{\partial F_2(\theta_2|\theta_1)}{\partial\theta_1} \leq 0$ and $\frac{dx_2^*}{d\theta_1} \leq 0$ are sufficient to ensure the third condition.

[10] Consider the following extension of our earlier parametric example. Let $f_1(\theta_1)$ be uniform over the interval $[\underline{\theta}_1,\bar{\theta}_1]$ so that $H(\theta_1) = \theta_1 - \underline{\theta}_1$. Now let $f_2(\theta_2|\theta_1)$ be exponentially distributed with mean θ_1 so that $F_2(\theta_2|\theta_1) = 1 - e^{-\frac{\theta_2}{\theta_1}}$ for $\theta_2 \in [0,\infty)$. We have $K(\theta_1,\theta_2) = \theta_2/\theta_1$. For a θ_1-type BU manager, a lower $\underline{\theta}_1$ implies a higher $H(\theta_1)$. Hence, we have more slack in both the first and second periods.

H3: Organizational slack is positively serially correlated.

III. EMPIRICAL METHODS

Sample and Data

The data for our study consists of the survey responses of both the *manager* and the *controller* of 104 BUs of six firms listed on the Amsterdam Exchanges with sales exceeding €2 billion. We describe the survey design and administration in detail in Appendix B. We define a business unit of a firm as an entity that (1) has its own general manager (hereafter, BU manager), (2) has a chief financial officer (hereafter, BU controller), and (3) is held responsible for a clearly defined part of the business by means of a periodic budget. The number of BUs from each firm range from a low of 9 to a high of 26 and sales of the median BU is about €150 million.

Measurement of Variables

Organizational Slack

As discussed in the introduction, we use perceived achievability of performance targets as an indicator of organizational slack. This is based on the intuition that easily achievable targets do not sufficiently motivate BU managers to operate at their full potential (or close to first-best levels) and allow them to earn rents (see also, Merchant and Manzoni 1989; Van der Stede 2000). Specifically, our survey instrument asks BU managers to describe their current performance targets on a five-point scale ranging from "easy to attain" to "practically unattainable" (see "m1" in Appendix B). Although the use of a single indicator to measure a complex construct is almost certainly associated with measurement error, the existing empirical literature has not developed a more comprehensive approach to measuring organizational slack.

To provide some support for reliability of *SLACK*, we follow Nunnally and Bernstein (1994) and examine temporal stability of the construct. Specifically, about a year after the original survey, a small subset of our sample BU managers were asked to reassess the achievability of their budgetary targets pertaining to the original survey period. We find a positive correlation ($\rho = 0.57$, $p = 0.11$) between *SLACK* and scores on this similar instrument in the follow-up survey. To provide support for external validity of *SLACK*, we correlate it with an indicator of BU managers' "budgetary participation" introduced by Hofstede (1967). Prior studies (Young 1985; Dunk 1993) suggest that budgetary participation is positively correlated with organizational slack. For a subset of BU managers with the available information, we find, as expected, that *SLACK* is low when budgetary participation is low and corporate headquarters set budgetary targets largely in a top-down manner ($\rho = 0.22$, $p = 0.03$).

As a proxy for a prior period's organizational slack (needed to test H3), we use a measure of the ease with which a BU achieved its past performance target (*BUDGET-1*). This measure asks controllers to indicate on a seven-point scale how well the BU performed relative to its performance target in the prior year ("c1" in Appendix B). *BUDGET-2* and *BUDGET-3* are similar measures of performance two and three years earlier, respectively. We use *BUDGET-1*, *BUDGET-2*, and *BUDGET-3* as past indicators of organizational slack.

Nature of Management Accounting Systems

Hypothesis 2 predicts that organizational slack is higher in settings where the managerial accounting system emphasizes decision-making information (higher m in our model) and generates fewer pieces of information to control the BU manager (lower n in our model). Since our empirical proxy for a management accounting system is the institution of the BU controller (as discussed in the introduction), we view the generation of more decision-making information and less control information as analogous to BU controllers focusing relatively more on their local as opposed to corporate responsibilities. Thus, we construct a measure of "controller BU focus," defined as the extent to which BU controllers focus relatively more on their local duties as members of the local management team as opposed to their corporate control responsibilities. We use controller BU focus as the empirical counterpart of a relatively high m (low n) in our model.

Specifically, our measure of controller BU focus (*CBF*) is based on 18 questionnaire items describing: (1) BU managers' influence on decisions regarding appointment and compensation of BU controllers (m2–m5), (2) corporate (or group) controllers' influence over the work of BU controllers (c2, c3), and (3) local authority to make internal BU specific accounting choices (e.g., local cost allocations, transfer pricing, valuation techniques, internal performance measurement, design of planning and budgeting systems) independently of corporate headquarters (c4–c15). Appendix C describes how we construct the *CBF* measure and provides detailed evidence on its reliability. To provide some support for external validity of this construct, we correlate *CBF* with (1) controllers' tenure in their BU (see our control variable *TIME* below) and (2) BU managers' operational autonomy from corporate headquarters (described in Section IV); two constructs which our field interviews suggest should be positively correlated with controller BU focus (see also, Simon et al. 1954). Using all available survey data, we find as expected that *CBF* is positively correlated with controllers' tenure ($\rho = 0.18$, $p = 0.04$) and BU managers' operational autonomy from corporate headquarters ($\rho = 0.49$, $p = 0.00$).

Extent of Information Asymmetry

We consider two characteristics of the BU environment as potential proxies for the extent of information asymmetry between corporate headquarters and BU management. We expect that, for rapidly growing BUs and/or for BUs operating in highly uncertain environments, corporate headquarters have more difficulty keeping track of internal BU developments and consequently, there is more information asymmetry (see also, Baiman et al. 1995; Nagar 2002; Christie et al. 2003). The first proxy we use is an indicator of BU growth (*GROWTH*) first introduced by Gupta and Govindarajan (1984). In particular, this survey item asks respondents to indicate the percentage of total sales for which the BU strategy is "increase sales and market share, be willing to accept low returns on investment in the short-to-medium term, if necessary." To reduce deviations from normality, we calculate the square root of the reported percentages. The second proxy we use is an indicator of environmental uncertainty (*ENVIRON*) akin to a measure employed by Gul and Chia (1994). In particular, this survey item asks BU managers and controllers to indicate the predictability of BUs' economic environment. Respondents describe on 1–7 Likert scales ranging from "highly predictable" to "highly unpredictable" the following six items: competitors' actions, market demands, production technology, product attributes/design, purchasing of supplies, and government regulation.

Other Variables

We also use two control variables that our field observations suggest may be related to both organizational slack and controller BU focus: *SIZE* is the natural logarithm (log) of the number of people employed in a BU; *TIME* is the log of the number of years the controller has held the position in the BU.

Descriptive Evidence

Table 1 reports descriptive statistics for our dependent and independent variables. On average, BU managers perceive their current targets to be achievable with high effort (mean of *SLACK* is 3.41), and prior period performance of their BU is slightly above target (mean of *BUDGET-1* is 4.18). The raw scores on *CBF* (average of all items used for calculating factor scores as described in Appendix C) range from 1.6 to 5.1 suggesting that controllers' BU focus varies greatly across BUs. The *GROWTH* variable implies that, on average, about 25 percent of BU sales are classified as generating low returns in the short-to-medium term. Also, a typical BU in our sample has 1,260 employees (*SIZE*), and, on average, a controller's tenure in a BU is approximately four years (mean of *TIME* is 4.21). Finally, we note that, with the exception of *SIZE*, most of the variation in the variables reported in Table 1 is at the BU rather than at the firm level (see the last two columns of Table 1).

To provide some supporting evidence that annual performance targets (and hence their achievability) play an important role in motivating BU managers, we note from our field observations that the primary incentive component of BU managers' compensation is an annual bonus. BU managers in our sample generally do not receive equity-based compensation. The median bonus is 30 percent, while the maximum is 65 percent of total compensation (untabulated information available for a subset of 95 BUs). On average, about 60 percent of the annual bonus is paid out based on BU performance relative to its target, about 20 percent is based on business group or firm results, about 15 percent based on nonfinancial measures and the rest is determined subjectively. Typically, BU targets that underlie annual bonuses are negotiated between the BU manager and a higher-level executive at the business group or corporate level. For example, 44 percent of BU managers in

TABLE 1
Descriptive Statistics[a]

Variable	n	Mean	Std. Dev.	Min	Max	ρ^c	F-test
SLACK	104	3.41	0.60	2.0	5.0	.039	.143
GROWTH	104	24.88	25.06	0.00	100.0	.087	.029
ENVIRON	104	3.29	0.56	1.8	4.8	.000	
CBF[b]	104	3.22	0.80	1.6	5.1	.049	.105
BUDGET-1	104	4.18	1.95	1.0	7.0	.000	
SIZE	104	1,260	3,948	9.0	39,000	.342	.000
TIME	104	4.21	5.14	0.0	33.0	.026	.215

[a] Statistics describe data before transformations (in subsequent analyses we use the square root of *GROWTH* and log of *SIZE* and *TIME*) and reverse coding (*SLACK*, *CBF*).

[b] We present descriptive statistics on scores calculated as averages of the 15 items aggregated in *CBF* (rather than less informative descriptive statistics on factor scores actually used in our regressions; see Appendix C).

[c] ρ estimates the proportion of variance accounted for by firm-level effects. It is equal to the correlation between variable scores of two randomly drawn BUs within the same firm (Snijders and Bosker 1999). $\rho = .000$ reported when the estimate is negative. The last column reports p-values associated with F-statistics testing the null hypothesis that there is no firm level effect.

SLACK = organizational slack (perceived achievability of a budgetary target);
GROWTH = "sales of products with a growth strategy" as a percentage of total sales;
ENVIRON = perceived predictability of the BU environment;
CBF = a measure of the extent to which BU controllers focus on their local responsibilities;
BUDGET-1 = BU's prior year's performance relative to its target;
SIZE = number of employees; and
TIME = number of years controller has held the position in a BU.

our sample describe the process of setting a budgetary target as "proposed by me, decided jointly," while about 40 percent report that higher-level management is more closely involved.

Our evidence also suggests that BU controllers' compensation plans differ from those of BU managers.[11] For instance, the explicit incentives based on a BU budgetary target are much weaker for a controller than for a manager; the median BU controller bonus is 15 percent of annual compensation, while the maximum is 20 percent. Instead, our field observations suggest that job security and promotions within the firm are a more important source of incentives for BU controllers than an annual bonus.

We conclude our presentation of descriptive evidence by examining univariate correlations among the variables of interest (Table 2). Notably *SLACK*, our dependent variable, is positively correlated with *CBF* and *BUDGET-1*, consistent with H2 and H3. Untabulated results further show no significant relation ($\rho = 0.13$, $p = 0.18$) between *SLACK* and the percentage of bonus based on a BU budgetary target and a significantly positive relation between *SLACK* and BU managers' participation in setting BU budgetary targets (as discussed in the measurement section, the latter relation provides support for external validity of the SLACK measure). Finally, there is some correlation among the independent variables, but none that suggests serious collinearity concerns.

Empirical Model

We test our hypotheses about the determinants of organizational slack via the following recursive model:

$$SLACK_{ij} = \delta_{0j} + \delta_1 GROWTH_{ij} + \delta_2 ENVIRON_{ij} + \delta_3 CBF_{ij} + \delta_4 BUDGET\text{-}1_{ij} + \delta_5 SIZE_{ij} + \delta_6 TIME_{ij} + \varepsilon_{ij}, \quad (9)$$

TABLE 2

Pearson Correlations among Dependent and Independent Variables
(n = 104)

	SLACK	*GROWTH*	*ENVIRON*	*CBF*	*BUDGET-1*	*SIZE*
GROWTH	0.170					
ENVIRON	−0.141	−0.025				
CBF	0.205*	−0.022	−0.255**			
BUDGET-1	0.206*	−0.220*	−0.069	−0.028		
SIZE	−0.110	−0.190	−0.007	0.045	0.117	
TIME	−0.038	−0.236*	−0.021	0.027	−0.006	0.167

**, * Correlations are significant at the 0.01 and 0.05 levels, respectively (two-tailed).
SLACK = organizational slack (perceived achievability of a budgetary target);
GROWTH = "sales of products with a growth strategy" as a percentage of total sales;
ENVIRON = perceived predictability of the BU environment;
CBF = a measure of the extent to which BU controllers focus on their local responsibilities;
BUDGET-1 = BU's prior year's performance relative to its target;
SIZE = number of employees; and
TIME = number of years controller has held the position in a BU.

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[11] Our evidence concerning BU controllers' compensation plans is admittedly limited. It comes from a follow-up survey of the annual bonus plans of 17 BU controllers who were participants in the original survey.

$$CBF_{ij} = \beta_{0j} + \beta_1 GROWTH_{ij} + \beta_2 ENVIRON_{ij} + \beta_3 SIZE_{ij} + \beta_4 TIME_{ij} + \varepsilon_{ij} \quad (10)$$

where the subscript i denotes a BU-level observation, the subscript $j = 1,...,6$ denotes firms. Hypothesis 1 corresponds to positive predictions about the δ_1 and δ_2 coefficients, while H2 and H3 correspond to positive predictions about δ_3 and δ_4 respectively.

Our recursive specification assumes that a controller's BU focus is determined by some of the same characteristics that determine organizational slack, but organizational slack itself is unlikely to influence the choice of *CBF*. That is, we assume *CBF* is effectively predetermined when firms are setting annual BU performance targets, and thus, *CBF* is uncorrelated with the error terms in Equation (9). Nonetheless as a robustness check, we also estimate Equation (9) using a two-stage estimation procedure that does not assume *CBF* is a predetermined variable (see Section IV).

Our specification also reflects that BU observations are not independent but clustered within six firms; hence, we include firm-specific intercepts (δ_{0j} and β_{0j}), estimate the system of Equations (9) and (10) by weighted least squares (WLS), which allows for firm-specific error terms, and adjust the standard errors for data clustering.[12]

IV. EMPIRICAL RESULTS

Hypotheses Tests

Table 3 presents our results: Panel A reports the results for Equation (9), while Panel B reports the results for Equation (10). Panel A shows that there is a significant positive relation ($p < 0.05$) between organizational slack (*SLACK*) and BU growth (*GROWTH*), our first proxy for information asymmetry between corporate headquarters and BU managers. This suggests that performance targets in rapidly growing BUs are easier to achieve and allow for more organizational slack than targets in more stable BUs, which is consistent with H1. However, we find no significant relation between organizational slack and environmental uncertainty (*ENVIRON*), our second proxy for information asymmetry. With the benefit of hindsight, we conjecture that our measure of environmental uncertainty is more a reflection of overall uncertainty about BU operations than a reflection of BU management's informational advantage over corporate headquarters.

Panel A of Table 3 further shows a significant positive relation ($p < 0.01$) between organizational slack (*SLACK*) and controller BU focus (*CBF*). Consistent with H2, this suggests that organizational slack is greater when BU controllers focus relatively more on providing information for decision making than for corporate control.

Panel A of Table 3 also reports a significant positive relation ($p < 0.01$) between organizational slack (*SLACK*) and prior period's slack as measured by BUs' prior period performance relative to budget (*BUDGET-1*). Consistent with H3, this suggests that organizational slack is positively serially correlated. That is, BU managers who have easily surpassed their performance targets in the past also expect to easily surpass their future performance targets. This finding is consistent with Indjejikian and Nanda (2002) and suggests that firms do not update BU managers' performance targets substantively to reflect

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[12] WLS, also referred to as feasible generalized least squares (Greene 2000), is an iterative procedure that obtains consistent estimates of firm-specific error variances and uses them in a subsequent step as weights to calculate coefficient estimates. For recursive systems of equations, WLS produces consistent coefficient estimates even in the presence of endogenous right-hand-side variables.

TABLE 3
Weighted Least Squares Estimates of the Determinants of Organizational Slack

$$SLACK_{ij} = \delta_{0j} + \delta_1 GROWTH_{ij} + \delta_2 ENVIRON_{ij} + \delta_3 CBF_{ij} + \delta_4 BUDGET\text{-}1_{ij} + \delta_5 SIZE_{ij} + \delta_6 TIME_{ij} + \varepsilon_{ij}$$

$$CBF_{ij} = \beta_{0j} + \beta_1 GROWTH_{ij} + \beta_2 ENVIRON_{ij} + \beta_3 SIZE_{ij} + \beta_4 TIME_{ij} + \varepsilon_{ij}$$

		Panel A: *SLACK*		Panel B: *CBF*	
Regressors[a]	**Predicted Sign**	**Coefficient**	**p-value**[b]	**Coefficient**	**p-value**[b]
GROWTH	+	0.050	.046	0.020	.890
ENVIRON	+	−0.073	.218	−1.301	.002
CBF	+	0.048	.000		
BUDGET-1	+	0.100	.000		
SIZE		−0.019	.372	0.413	.302
TIME		0.014	.715	0.083	.647
Adjusted R^2		.183		.065	
n		104		104	

[a] Firm-specific intercepts are not reported.
[b] p-values are two-tailed, based on standard errors robust to clustering of data.

SLACK = organizational slack (perceived achievability of a budgetary target);
GROWTH = "sales of products with a growth strategy" as a percentage of total sales;
ENVIRON = perceived predictability of the BU environment;
CBF = a measure of the extent to which BU controllers focus on their local responsibilities;
BUDGET-1 = BU's prior year's performance relative to its target;
SIZE = number of employees; and
TIME = number of years controller has held the position in a BU.

past performance, because otherwise the achievability of performance targets in the future would be independent of past performance.[13]

Finally, Panel B of Table 3 suggests that controller BU focus (*CBF*) is significantly negatively related to environmental uncertainty or *ENVIRON* ($p < 0.01$) but unrelated to *GROWTH*, *SIZE*, and *TIME*. This implies that, in more uncertain environments, BU controllers tend to focus relatively more on their control responsibilities than on the provision of decision-making information. We rely on this finding in Section IV when we obtain weighted two-stage least squares estimates of Equation (9) using *ENVIRON* as an instrumental variable.

Additional Evidence

Time-Series Correlation in Organizational Slack

The evidence presented in Table 3 is based on estimates of the relations between organizational slack and its (cross-sectional) determinants of information asymmetry, controller BU focus, and past BU performance relative to target. In this section, we present additional support for H3 based on time-series evidence of how BU performance relative to target is serially correlated over a three-year period. Panel A of Table 4 shows that about

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[13] In a related study, Leone and Rock (2002) find that managers' performance targets are revised upward, following superior past performance. To the extent that their finding is descriptive of practice, it suggests that the serial correlation in perceived achievability of performance targets is likely to be muted.

TABLE 4
Time-Series Evidence on Achievability of Performance Targets

Panel A: Actual Performance Relative to Budgetary Target in Last Three Years

	n	Below Target (<4)	Same as Target (=4)	Above Target (>4)
BUDGET-1	104	40 (38.5%)	21 (20.2%)	43 (41.3%)
BUDGET-2	98	39 (39.8%)	27 (27.6%)	32 (32.7%)
BUDGET-3	93	27 (29.0%)	31 (33.3%)	35 (37.6%)

Panel B: Actual Performance Relative to Budgetary Target Conditional on Performance Relative to Budgetary Target in Prior Year

		Year t		
Year $t-1$	n	Below Target (<4)	Same as Target (=4)	Above Target (>4)
Below Target (<4)	66	43 (65.2%)	10 (15.2%)	13 (19.7%)
Above Target (>4)	67	14 (20.9%)	8 (11.9%)	45 (67.2%)
			χ^2-test: 32.63 (p = 0.000)[a]	

Panel C: Actual Performance Relative to Budgetary Target Conditional on Performance Relative to Budgetary Target in Two Prior Years

		BUDGET-1		
	n	Below Target (<4)	Same as Target (=4)	Above Target (>4)
BUDGET-2 < 4 or *BUDGET-3* < 4	44	27 (61.4%)	6 (13.6%)	11 (25.0%)
BUDGET-2 ≥ 4 and *BUDGET-3* ≥ 4	22	4 (18.2%)	2 (9.1%)	16 (72.7%)
			χ^2-test: 14.24 (p = 0.001)[a]	

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[a] χ^2-statistic tests the null hypothesis that the two classifications are independent.
BUDGET-t = BU performance relative to target t years ago (t = 1,2,3).

62 percent of BUs met or exceeded their budgetary target in the prior year. The proportion of BUs that met or exceeded their target two and three years earlier are 60 percent and 71 percent, respectively.

Panels B and C of Table 4 describe the serial correlations in these proportions. Panel B shows that about 79 percent of BUs that met or exceeded their target in a given year also met or exceeded their target in the following year, while only 35 percent of BUs that did not meet their target in a year, succeeded in the following year (χ^2-test rejects the null hypothesis of zero time-series correlation with $p < 0.01$). Panel C shows that 82 percent of BUs that met or exceeded their performance targets in two consecutive years also met or exceeded their target in the subsequent year. Again, the null hypothesis of zero time-series correlation is rejected ($p < 0.01$).

Value of Decision-Making Information

An implicit feature of our framework in Section II is that a management accounting system that provides more decision-making information is valuable to the firm, even though it may also increase organizational slack. Since controller BU focus is our proxy for the relative extent to which management accounting systems emphasize decision-making information, an implicit assumption in our empirical analysis is that a BU controller's focus improves local decision making. While we cannot directly measure improvements in local decision making, we provide some validation for this implicit assumption by constructing a measure (*SATIS*) of how satisfied BU managers are with the decision support provided by their controllers.

Specifically, we construct *SATIS* based on two questions. The first question asks BU managers to evaluate the services provided by their BU controllers along three dimensions (m6–m8) that typically describe the main tasks of a controller's department. The second question asks BU managers about their access to information potentially available from their management accounting systems, e.g., product profitability, externally oriented information (see also, Chenhall and Morris 1986). BU managers indicate on two six-point Likert scales (m9–m14) the extent to which the listed information items are (1) important, and (2) available. We calculate *SATIS* as the average ranking on items m6–m8 and the difference between (1) and (2).[14]

We test the relation between *SATIS* and controller BU focus (*CBF*) akin to our earlier specification of the relation between organizational slack and *CBF* (i.e., Equation (9)). Table 5 presents the results of our WLS estimation. As expected, we find that controller BU focus (*CBF*) is positively associated with *SATIS* ($p = 0.01$), which provides some support for our implicit assumption that greater controller BU focus can be value enhancing. However, we

TABLE 5
Weighted Least Squares Estimates of the Determinants of BU Managers' Satisfaction

$$SATIS_{ij} = \gamma_{0j} + \gamma_1 GROWTH_{ij} + \gamma_2 ENVIRON_{ij} + \gamma_3 CBF_{ij} + \gamma_4 BUDGET\text{-}1_{ij} + \gamma_5 SIZE_{ij} + \gamma_6 TIME_{ij} + \varepsilon_{ij}$$

		WLS Estimates *SATIS*	
Regressors[a]	**Predicted Sign**	**Coefficient**	**p-value**[b]
GROWTH		−2.952	.001
ENVIRON		0.838	.721
CBF	+	1.532	.010
BUDGET-1		1.092	.263
SIZE		−3.990	.153
TIME		4.397	.000
Adjusted R^2		.174	
n		104	

[a] Firm-specific intercepts are not reported.
[b] p-values are two-tailed, based on standard errors robust to clustering of data.
SATIS = managers' satisfaction with information provided by their controlling department,
GROWTH = "sales of products with a growth strategy" as a percentage of total sales;
ENVIRON = perceived predictability of the BU environment;
CBF = a measure of the extent to which BU controllers focus on their local responsibilities;
BUDGET-1 = BU's prior year's performance relative to its target;
SIZE = number of employees; and
TIME = number of years controller has held the position in a BU.

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[14] The data support unidimensionality of the four-item construct; Cronbach's α is 0.80. As an empirical matter, the difference between importance and availability is negative for 11 BUs. Since a negative difference has no obvious interpretation, we replace negative values with zero. We also consider using the absolute value of the differences as well as leaving the differences negative. Our results remain qualitatively unchanged.

acknowledge that the first survey question that comprises the *SATIS* measure may reflect not only BU managers' satisfaction with decision support of their controllers, but also BU managers' satisfaction with informational rents they garner via their controllers' efforts.

We also find that BU managers' satisfaction is negatively associated with BU growth ($p < 0.01$) and positively associated with BU controllers' experience in the BU ($p < 0.01$). The former finding is consistent with prior literature (e.g., Sathe 1982; Granlund and Lukka 1998b) that suggests that controllers can support day-to-day or operational decisions better than strategic decisions (the latter type of decisions is presumably more important for growing BUs).

Sensitivity Analysis

Alternative Measures of Controller BU Focus

As described in Section III and Appendix C, we construct our measure of controller BU focus (*CBF*) from 18 survey items. In this section, we examine the extent to which our findings are sensitive to different ways of measuring *CBF*. First, we reproduce the approach in Appendix C, but exclude four of the 18 survey items (m2–m5) that were answered by BU managers rather than BU controllers. Second, instead of the approach employed in Appendix C, we measure *CBF* as an equally weighted average of the 18 survey items. Alternatively, we calculate the average of the BU ranks for (rather than an equally weighted average of) the six underlying indicators of controller BU focus detailed in Appendix C. Employing these three alternative measures in our tests does not materially affect our results.

We also address the possibility that *CBF* is spuriously associated with *SLACK* because *CBF* also reflects BU managers' decision rights to operate autonomously from corporate headquarters. To address this concern, we construct a measure of BU managers' operational autonomy (*DECENTR*) based on the average responses to 16 survey questions (using 1–7 Likert scales) concerning BU managers' decision rights or responsibilities in four areas—marketing (four questions), financial (five questions), operational (five questions), and purchasing (two questions). We construct an alternative measure of controller BU focus as the residual from a WLS regression of *CBF* on *DECENTR*. Untabulated results reveal that the positive association between controller BU focus and organizational slack remains significant ($p = 0.02$) even when we use this alternative measure.

Alternative Estimation Techniques

To confirm the robustness of our estimation techniques, we consider two alternative approaches. First, we estimate an ordered logit model of organizational slack described in Equation (9) and find qualitatively similar results (untabulated) as in Table 3. Second, we estimate Equation (9) via a weighted two-stage least squares (2SLS) approach where we use firm-specific instruments $ENVIRON_j$ in the first stage. Consistently with the results presented in Table 3, the results of the 2SLS approach (Table 6) suggest that organizational slack is positively associated with BU growth ($p = 0.02$), controller BU focus ($p < 0.01$), and BUs' prior period performance relative to budget ($p < 0.01$). That said, we acknowledge that our 2SLS results should be interpreted with caution since the 2SLS procedure entails a reduction in the degrees of freedom that is particularly critical for our study given our limited sample size.

V. SUMMARY AND CONCLUSIONS

We examine the determinants of organizational slack for a sample of 104 business unit (BU) managers of large decentralized firms. To guide our study, we characterize the relationship between our sample firms and their better-informed BU managers as a standard

TABLE 6
Weighted Two-Stage Least Squares Estimates of the Determinants of Organizational Slack

$$SLACK_{ij} = \delta_{0j} + \delta_1 GROWTH_{ij} + \delta_3 CBF_{ij} + \delta_4 BUDGET\text{-}1_{ij} + \delta_5 SIZE_{ij} + \delta_6 TIME_{ij} + \varepsilon_{ij}$$

Regressors[a]	Predicted Sign	SLACK[c] Coefficient	p-value[b]
GROWTH	+	0.052	.018
CBF	+	0.074	.001
BUDGET-1	+	0.104	.000
SIZE		−0.030	.386
TIME		0.012	.779
Adjusted R^2		.175	
n		104	

[a] Firm-specific intercepts are not reported.
[b] p-values are two-tailed, based on standard errors robust to clustering of data.
[c] The instruments used are $GROWTH_j$, $ENVIRON_j$, $BUDGET\text{-}1_j$, $SIZE_j$, and $TIME_j$.

SLACK = organizational slack (perceived achievability of a budgetary target);
GROWTH = "sales of products with a growth strategy" as a percentage of total sales;
ENVIRON = perceived predictability of the BU environment;
CBF = a measure of the extent to which BU controllers focus on their local responsibilities;
BUDGET-1 = BU's prior year's performance relative to its target;
SIZE = number of employees; and
TIME = number of years controller has held the position in a BU.

adverse selection problem that gives rise to organizational slack represented by easily achievable performance targets. Consistent with this characterization, we find that the extent of information asymmetry between corporate headquarters and BU managers (measured by the superior local knowledge implicit in growing BUs) is positively associated with organizational slack.

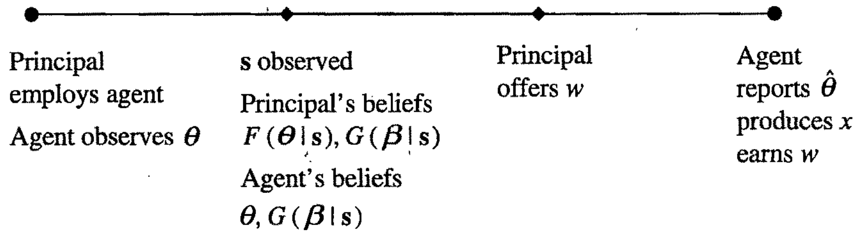
We characterize a management accounting system (represented by a BU controller) as providing decision-making information to local managers while mitigating control problems between corporate headquarters and local management. We find that organizational slack is higher when BU controllers focus relatively more on their local decision-support responsibilities than on their control responsibilities. At the same time, we find some evidence that a greater local focus by BU controllers is indeed beneficial. Taken together, these findings imply that management accounting institutions designed to improve local decision making can exacerbate control problems in decentralized firms. Conversely, management accounting practices that emphasize corporate control can undermine local decision making to the detriment of the firm.

We also find evidence that organizational slack is positively serially correlated. In particular, we find that BU managers' tendencies to meet or exceed their performance targets are persistent; moreover, past successes do not seem to translate to harder-to-achieve targets in the future. We interpret this evidence as largely consistent with theoretical arguments that, in ongoing firm-manager relationships, firms' explicit or implicit commitment to discount managers' past performance in setting their future performance targets improves managerial incentives to increase firm value.

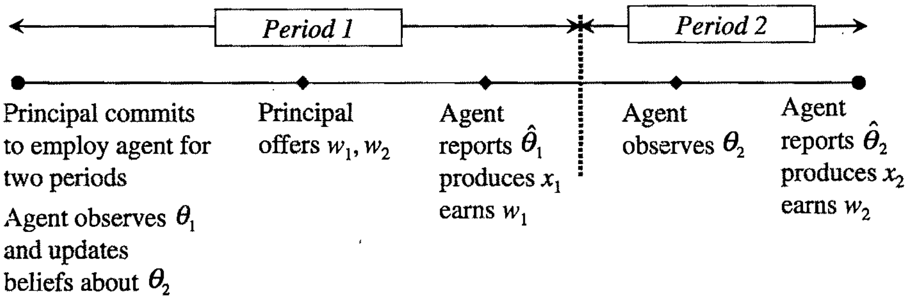
Our study is subject to a number of limitations, most of which we acknowledge and address in our study. One important limitation is that, due to the lack of archival data on firms' internal accounting systems, we rely on BU managers' perceptions of the achievability of their performance targets to measure organizational slack. Although we acknowledge the presence of measurement error inherent in such an approach, our tests of construct validity and other sensitivity checks provide some reassurance that measurement error is limited. Perhaps the most important limitation of our study is its small sample of participating organizations (104 BUs of six firms with corporate headquarters in The Netherlands), which likely reduces the generalizability of our findings. Documenting the role of accounting systems and institutions in a broader sample of decentralized organizations would clearly be valuable.

APPENDIX A

Timeline 1: One-Period Adverse Selection Model


Principal employs agent
Agent observes θ
s observed
Principal's beliefs
$F(\theta \mid s), G(\beta \mid s)$
Agent's beliefs
$\theta, G(\beta \mid s)$
Principal offers w
Agent reports $\hat{\theta}$ produces x earns w

Timeline 2: Two-Period Extension


Period 1
Period 2
Principal commits to employ agent for two periods
Agent observes θ_1 and updates beliefs about θ_2
Principal offers w_1, w_2
Agent reports $\hat{\theta}_1$ produces x_1 earns w_1
Agent observes θ_2
Agent reports $\hat{\theta}_2$ produces x_2 earns w_2

APPENDIX B

Survey Administration

We invited 19 firms that met our selection criteria (see Section III) to participate in the research project. Seven firms agreed to give access to their business units (BUs) and to provide all required information. Sales of these firms ranged between €2 and 15 billion, their primary activities were food processing (one firm), manufacturing and trading (four firms), and services (two firms).

In each firm we interviewed five to ten controllers at different organizational levels (48 interviews in total) and studied internal documents (accounting manuals, organizational charts). We sent 363 (eight-page) questionnaires to managers and controllers of 188 BUs of the seven firms between February 2000 and November 2001.[15] The first mailing went from the headquarters and included a recommendation letter from the corporate controlling director. There were two follow-up mailings. In June 2002, we mailed participating controllers and managers the final research report and invited them to fill in an abbreviated (two-page) version of the original questionnaire. There was no follow-up mailing (in order to honor an agreement with managers sponsoring the study).

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[15] We followed recommendations of the Total Design Method (Dillman 1978). There was a separate survey for each of the firms. The number of selected BUs varied between 13 and 33 per firm. Most of excluded BUs were relatively small, were about to be divested, were undergoing a reorganization and/or had recent changes in the top two managerial positions. We slightly modified the questionnaires in each of the separate surveys to make them applicable in different contexts and to improve measurement. These modifications mostly concerned control variables.

We received 308 usable questionnaires (a response rate of 85 percent) from 178 BUs (95 percent of BUs contacted). Data on organizational slack is available in 107 BUs of six firms. After excluding three BUs with missing data on BUDGET-1 we obtain the final sample of 104 BUs.

Questionnaire Items[16]

Generally, the budget targets are: m1

Scale: 1 easy to attain
2 attainable with normal effort
3 attainable with high effort
4 attainable with very high effort
5 practically unattainable

How is authority for the following decisions divided between you, as the general manager of the business unit, and higher level controlling?

Scale: 1 Decision is taken at our business unit without consulting higher levels.
7 Decision is taken at higher levels without consulting our business unit.

Hiring of the BU controller	m2
Transfer of the BU controller	m3
Salary increase of the BU controller	m4

How is authority divided between your business unit and some higher level for each of the following classes of decisions?

Scale: 1 Decision is taken at our business unit without consulting higher levels.
7 Decision is taken at higher levels without consulting our business unit.

The hiring of business unit controlling staff m5

What was the performance of your business unit in the last three years, relative to your budget and relative to the overall firm performance? c1

1 far below the budget	1 far below the overall firm performance
4 about the same as the budget	4 about the same as the overall firm performance
7 far above the budget	7 far above the overall firm performance

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[16] Items answered by managers are coded m1–m14, items answered by controllers c1–15.

	2001	2000	1999
Compared to the budget			
Compared to the overall firm performance			

Comparing the general manager of your business unit and the business group controller, who of them assigns relatively more tasks to your controlling department? c2

Scale: 1 Almost all tasks come from the general manager.
4 Both assign about the same amount of tasks.
7 Almost all tasks come from the group controller.

Who of them typically has more influence on final decisions when their views differ? c3

Scale: 1 The general manager makes almost always the final decision.
4 The general manager makes the final decision in some areas, the group controller in others.
7 The group controller makes almost always the final decision.

How is authority divided between your business unit (BU) and higher levels when a decision to change the following accounting techniques and reporting procedures is made?

Scale: 1 The decision about the change is taken at our BU without consulting higher levels.
6 Decision about the change is taken at higher levels without consulting our BU.

Allocation of manufacturing overhead	c4
Allocation of marketing costs	c5
Transfer prices for transactions within your BU	c6
Expensing versus capitalizing costs (e.g., provisions, writing-off receivables)	c7
Inventory valuation	c8
Financial performance indicators to be used within your business unit	c9
Nonfinancial performance indicators to be used within your BU	c10
Design of your internal reports (e.g., frequency, breadth of information)	c11
Budgeting process within your BU	c12
Short-range financial planning	c13
Strategic planning	c14
Criteria to evaluate capital investment (e.g., accounting measure, satisfactory rate)	c15

How would you evaluate the services provided by your controlling department along the following dimensions?

Scale: 1 I am satisfied with how they carry out this task.
7 I am not satisfied at all with how they carry out this task.

Reporting to you and business unit management m6

Assisting operating managers in current analysis and interpretation of accounting information m7

Providing accounting information for solving business problems on a special studies basis m8

How important are the following information items in managing your business unit?
To what extent are they actually available to managers?

Scale: 1 Not important/Not available to managers in our business unit.
7 Extremely important/Routinely provided to managers in our business unit.

Externally oriented information (on customers, competitors, etc.)	m9
Forward-looking information (forecasts, scenario analysis, etc.)	m10
Nonfinancial information (employee turnover, efficiency, etc.)	m11
Frequent reports (e.g., daily, weekly)	m12
Precise targets for all departments within your business unit	m13
Product profitability information	m14

APPENDIX C

Constructing a Measure of Controller BU Focus

We draw on Simon et al. (1954) to identify two broad dimensions of controller BU focus: authority relations of controllers and the structure of the accounts and reports at the BU. We operationalize the two broad dimensions with 18 questionnaire items (m2–m5, c2–c15) based on the extant literature and extensive interviews with controllers.

We examine reliability of the questionnaire items as follows. We analyze factorial structure of the 18 questionnaire items. Given that theoretical guidance on the structure is limited, we derive a six-factor solution (explaining 75 percent of the variance) from the data examining the scree plot, eigenvalues, and relatedness of the items (several items ask similar questions) loading on the same factor. After dropping three items (c11, c14, c15), we achieve an acceptable fit with the following six factors: (1) a BU manager's autonomy to appoint and compensate his BU controller (m2–m5), (2) actual influence of the group or corporate controller on the work of the BU controller (c2, c3), (3) BU autonomy to change cost allocation and transfer pricing techniques (c4–c6), (4) autonomy to change valuation techniques (c7, c8), (5) autonomy to design internal performance indicators (c9, c10), and (6) autonomy to design local planning and budgeting systems (c12, c13). Reliability tests as measured by Cronbach's α, or Pearson correlation (r) in case of two items, are satisfactory: construct (1) $\alpha = 0.83$, (2) $r = 0.35$ ($p = 0.00$), (3) $\alpha = 0.80$, (4) $r = 0.58$ ($p = 0.00$), (5) $r = 0.67$ ($p = 0.00$), (6) $r = 0.45$ ($p = 0.00$). In addition, we also test temporal stability of selected items: construct (2) item c2, $r = -0.35$ ($p = 0.17$), (4) item c7, $r = 0.55$ ($p = 0.03$), (5) item c9, $r = 0.48$ ($p = 0.06$), (6) item c12, $r = 0.41$ ($p = 0.11$). Based on the factor analysis, we obtain six constructs calculated as equally weighted averages of (reverse-coded) item responses loading on the same factor.

Conceptually, the six constructs are first-order factors (Rindskopf and Rose 1988), which adjust the weighting of items that are similar. For example, the fact that we include four items regarding hiring and compensation of controllers, but only two items regarding autonomy to change valuation techniques, does not mean that the former has a greater weight than the latter in the overall measure of controller BU focus.

We conceptualize the measure of controller BU focus as a second-order factor with the six constructs as *formative* indicators (Diamantopoulos and Winklhofer 2001). This is in line with Simon et al. (1954, 20) who argue that more controller BU focus along one of its dimensions by no means implies that it is necessary or desirable to have more BU focus in other dimensions. Thus, the six constructs need not be correlated for our measure of controller BU focus to be valid (Bollen and Lennox 1991); unidimensionality and Cronbach's α cannot be used as criteria to establish validity as in the standard case with *reflective* indicators. We calculate controller BU focus, CBF, as an equally weighted average of the six constructs after standardization.

Finally, we acknowledge that there are alternative ways of aggregating information from the 18 items into a measure of controller BU focus. Section IV shows that our findings are not affected by calculating our measure of controller BU focus in different ways.

REFERENCES

Antle, R., and G. D. Eppen. 1985. Capital rationing and organizational slack in capital budgeting. *Management Science* 31 (2): 163–174.

———, and J. Fellingham. 1995. Information rents and preferences among information systems in a model of resource allocation. *Journal of Accounting Research* 33: 41–58.

Baiman, S., and J. Demski. 1980. Economically optimal performance evaluation and control systems. *Journal of Accounting Research* 18: 184–220.

———, and J. H. Evans III. 1983. Pre-decision information and participative management control systems. *Journal of Accounting Research* 21: 371–395.

———, D. F. Larcker, and M. V. Rajan. 1995. Organizational design for business units. *Journal of Accounting Research* 33: 205–229.

———, and M. Rajan. 1998. Optimal cost targets and incentives for acquiring expertise. *Review of Accounting Studies* 3: 231–259.

———, and K. Sivaramakrishnan. 1991. The value of private pre-decision information in a principal agent context. *The Accounting Review* 66: 747–766.

Baron, D. P., and D. Besanko. 1984. Regulation and information in a continuing relationship. *Information Economics and Policy* 1: 447–470.

———, and R. B. Myerson. 1982. Regulating a monopolist with unknown costs. *Econometrica* 50: 911–930.

Bollen, K. A., and R. Lennox. 1991. Conventional wisdom on measurement: A structural equation perspective. *Psychological Bulletin* 110: 305–314.

Chenhall, R. H., and D. Morris. 1986. The impact of structure, environment and interdependence on the perceived usefulness of management accounting systems. *The Accounting Review* 61 (1): 16–35.

Chow, C. W., J. C. Cooper, and W. S. Waller. 1988. Participative budgeting: Effects of a truth-inducing pay scheme and information asymmetry on slack and performance. *The Accounting Review* 63 (1): 111–122.

Christensen, J. 1981. Communication in agencies. *Bell Journal of Economics*: 661–676.

Christie, A. A., M. P. Joye, and R. L. Watts. 2003. Decentralization of the firm: Theory and evidence. *Journal of Corporate Finance* (9): 3–36.

Demski, J. S., and G. A. Feltham. 1976. *Cost Determination: A Conceptual Approach.* Ames, IA: Iowa State University Press.

Diamantopoulos, A., and H. M. Winklhofer. 2001. Index construction with formative indicators: An alternative to scale development. *Journal of Marketing Research* 38 (May): 269–277.

Dillman, D. A. 1978. *Mail and Telephone Surveys: The Total Design Method.* New York, NY: John Wiley.

Dunk, A. S. 1993. The effect of budget emphasis and information asymmetry on the relation between budgetary participation and slack. *The Accounting Review* 68 (2): 400–410.

Granlund, M., and K. Lukka. 1998a. It's a small world of management accounting practices. *Journal of Management Accounting Research* 10: 153–179.

———, and ———. 1998b. Towards increasing business orientation: Finnish management accountants in a changing cultural context. *Management Accounting Research* 9 (2): 185–211.

Greene, W. H. 2000. *Econometric Analysis*. 4th edition. Englewood Cliffs, NJ: Prentice Hall.

Gul, F. A., and Y. M. Chia. 1994. The effects of management accounting systems, perceived environmental uncertainty and decentralization on managerial performance: A test of three-way interaction. *Accounting, Organizations and Society* 19 (4–5): 413–426.

Gupta, A. K., and V. Govindarajan. 1984. Business unit strategy, managerial characteristics, and business unit effectiveness at strategy implementation. *Academy of Management Journal* 27 (1): 259–285.

Hofstede, G. H. 1967. *The Game of Budget Control.* Assen, The Netherlands: Van Gorcum.

Hopper, T. M. 1980. Role conflicts of management accountants and their position within organization structures. *Accounting, Organizations, and Society* 5 (4): 401–411.

Indjejikian, R., and D. Nanda. 2002. Executive target bonuses and what they imply about performance standards. *The Accounting Review* 77 (4): 793–819.

Institute of Management Accountants (IMA). 1999. *Counting More, Counting Less: Transformation in the Management Accounting Profession: The 1999 Practice Analysis of Management Accounting*. Montvale, NJ: IMA.

Jablonsky, S. F., and P. J. Keating. 1998. *Changing Roles of Financial Management: Integrating Strategy, Control, and Accountability.* Morristown, NJ: Financial Executives Research Foundation.

Laffont, J-J., and J. Tirole. 1986. Using cost observations to regulate firms. *Journal of Political Economy* 94 (3): 614–641.

———, and ———. 1993. *A Theory of Incentives in Procurement and Regulation.* Cambridge, MA: MIT Press.

Leone, A. J., and S. Rock. 2002. Empirical tests of budget ratcheting and its effects on managers' discretionary accrual choices. *Journal of Accounting and Economics* 33: 43–67.

Merchant, K. A. 1985. Budgeting and the propensity to create budgetary slack. *Accounting, Organizations and Society* 10 (2): 201–210.

———, and J-F. Manzoni. 1989. The achievability of budget targets in profit centers: A field study. *Accounting, Organizations and Society* 14 (3): 539–558.

Milgrom, P., and J. Roberts. 1992. *Economics, Organization and Management.* Englewood Cliffs, NJ: Prentice Hall.

Nagar, V. 2002. Delegation and incentive compensation. *The Accounting Review* 77 (April): 231–255.

Nunnally, J. C., and I. H. Bernstein. 1994. *Psychometric Theory.* Third edition. New York, NY: McGraw-Hill.

Onsi, M. 1973. Factor analysis of behavioral variables affecting budgetary slack. *The Accounting Review* 48 (3): 535–548.

Penno, M. 1984. Asymmetry of pre decision information and managerial accounting. *Journal of Accounting Research* 22: 177–191.

Rindskopf, D., and T. Rose. 1988. Some theory and applications of confirmatory second-order factor analysis. *Multivariate Behavioral Research* 23: 51–67.

Salanié, B. 1997. *The Economics of Contracts.* Cambridge, MA: The MIT Press

San Miguel, J. G., and V. Govindarajan. 1984. The contingent relationship between the controller and internal audit functions in large organizations. *Accounting, Organizations and Society* 9: 179–188.

Sathe, V. 1978. Who should control the division controller? *Harvard Business Review* 56: 99–104.

———. 1982. *Controller Involvement in Management.* Englewood Cliffs, NJ: Prentice Hall.

Simon, H. A., H. Guetskow, G. Kozmetsky, and G. Tyndall. 1954. *Centralization vs. Decentralization in Organizing the Controller's Department.* New York, NY: Controllership Foundation.

Snijders, T. A. B., and R. J. Bosker. 1999. *Multilevel Analysis: An Introduction to Basic and Advanced Multilevel Modeling.* London, U.K.: SAGE Publications.

Van der Stede, W. 2000. The relationship between two consequences of budgetary controls: Budgetary slack creation and managerial short-term orientation. *Accounting, Organizations and Society* 25 (6): 609–622.

Waller, W. S. 1988. Slack in participative budgeting: The joint effect of a truth-inducing pay scheme and risk preferences. *Accounting, Organizations and Society* 13 (January): 87–98.

Young, S. M. 1985. Participative budgeting: The effects of risk aversion and asymmetric information on budgetary slack. *Journal of Accounting Research* 23 (2): 829–842.

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The Effects of Joint Provision and Disclosure of Nonaudit Services on Audit Committee Members' Decisions and Investors' Preferences

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ABSTRACT: Recent corporate governance reforms that require audit committees to pre-approve audit and nonaudit services increase audit committees' accountability to third parties for actual auditor independence and audit quality. Other SEC reforms mandate the disclosure of fees for auditor-provided services and are aimed at influencing investors' perceptions of auditor independence. These fee disclosures also reveal audit committees' pre-approval decisions, enhancing public accountability. Thus, audit committees may be less willing to hire auditors for nonaudit services to avoid fee disclosures, even when joint provision improves audit quality. One hundred experienced corporate directors, responding as audit committee members or investors, participated in an experiment in which we manipulated the effect of the auditor's provision of nonaudit services on audit quality and the fee disclosure requirement. We find that audit committee members are more likely to recommend joint provision if audit quality improves, consistent with investors' preferences. However, unlike investors, committee members are more reluctant to recommend joint provision when public disclosures are required,

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even at the expense of audit quality. These findings offer evidence about an indirect effect of recent reforms.

Keywords: *audit committees; nonaudit services; disclosure; audit quality.*

Data Availability: *Contact Professor McDaniel.*

I. INTRODUCTION

In order to promote effective corporate governance, the Securities and Exchange Commission (SEC) and Congress (through the Sarbanes-Oxley Act of 2002, U.S. House of Representatives 2002) passed reforms aimed at increasing audit committees' accountability to investors for audit and financial reporting quality (see, Unger 2001; Turner 2002). Specifically, audit committees now pre-approve both audit and allowable nonaudit services.[1] The SEC acknowledged that certain nonaudit services may improve audit quality and, thus, recommended that audit committees consider the effects of nonaudit services on audit quality (Turner 2001). However, in making a pre-approval decision, the committee is charged with evaluating the compatibility of allowable nonaudit services (i.e., joint service provision) with the auditor remaining independent.

Complementing these regulations are additional disclosure mandates aimed at reversing the erosion of investor confidence in the quality of audits and capital markets (Turner 2001; Pitt 2002; Federal Register 2003). For example, public companies also disclose that the audit committee pre-approved all audit and auditor-provided nonaudit services and that it expressly considered the service compatibility with maintaining the auditor's independence (Unger 2001). A second mandate requires companies to report the total amount of audit and nonaudit service fees paid to the auditor. Together, these disclosures are designed to help investors assess auditor independence (audit quality) and improve investor confidence (SEC 2000).

The audit and nonaudit service fee disclosure requirement is expected to affect investor perceptions of auditor independence and their behaviors; however, these disclosures may also influence audit committees' decisions (behaviors) related to approving nonaudit services (Unger 2001). For example, by disallowing nonaudit services from the auditor, audit committees can avoid disclosures that might raise investor concerns about auditor independence (Abbott et al. 2003; Dopuch et al. 2003). However, if joint service provision improves audit quality, then disallowing it may not be in investors' best interests. Investigating the extent to which required disclosures might affect audit committee decisions, therefore, is important in evaluating the possible direct and indirect consequences of existing and proposed regulations.

We design an experiment to provide empirical evidence on three broad questions related to the consequences of these regulations on audit committees' decisions. First, do audit committees consider the nonaudit service effects on audit quality in deciding whether to approve the purchase of nonaudit services from the auditor? Second, to what extent do the mandated disclosures affect audit committees' pre-approval decisions? Finally, to what extent are audit committees' decisions regarding the auditors' joint service provision consistent with investors' preferences?

Participants in our experiment are 100 corporate directors who attended one of several KPMG Audit Committee Institute Roundtables during 2003. Eighty-one of these 100 directors assume roles of audit committee members and recommend whether a nonaudit service be jointly provided by the auditor or provided by an unaffiliated audit firm. In a 2 × 2 between-participants design, we manipulate the type of nonaudit service (risk management service versus human resource management service) and the type of company (public versus private). Because we cannot directly manipulate the *perceived* effect of the nonaudit service on audit quality (the variable of interest), we also instruct participants that the risk management (human resource management) nonaudit service will improve (have no effect on) audit quality. The type of company serves as a manipulation for whether there is/is not a requirement for public disclosures to investors (public versus private company, respectively).

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[1] The SEC prohibits certain nonaudit services that, if provided to an audit client, would impair an auditor's independence and, thus, reduce audit and financial reporting quality.

Related to our first two questions, we find that the pre-approval decision for joint service provision is influenced by whether the audit committee believes audit quality improves. We further find that committees are less likely to approve joint service provision when public disclosures are required, even if they believe joint provision results in a higher quality audit. While disclosures appear to increase some committees' sensitivities to auditor objectivity impairment, other committees pass up joint provision to avoid disclosures and potential investor concerns over auditor independence.[2]

To address the third question, the remaining 19 of the 100 directors assume roles as investors in the public company and are asked to indicate which firm they would prefer provide the risk management services. In addition to public disclosures related to audit and nonaudit service fees, investor participants receive voluntarily disclosed information that the joint service provision would improve audit quality. Our findings suggest that, relative to audit committee members, investors believing that the joint provision results in a higher audit quality are *more likely* to prefer the auditors provide the risk management services.[3] Together, our results suggest the possibility of unspecified effects of some recent auditor independence regulations, and also that additional disclosure related to nonaudit services may be warranted.

The paper proceeds as follows. Section II develops the research motivation and hypotheses for the effects of regulations on audit committee decisions. Sections III and IV then discuss the related experimental design and results, respectively. Section V presents the research question, design, and results comparing audit committee judgments and investor preferences. A final section discusses some implications of our findings and concludes the paper.

II. BACKGROUND AND HYPOTHESES DEVELOPMENT

Audit Committee Responsibility for Auditor Independence and Audit Quality

Enron and other recent notable accounting frauds illustrate the importance of increasing audit committee governance accountability particularly as it relates to ensuring auditor independence and audit quality. Among these new responsibilities, audit committees now hire and fire the auditor and are required to pre-approve nonaudit services. While these recent legislative reforms are required only for public companies, private companies and their audit committees are also affected. For example, one recent survey found that "30 percent of fast-growth private companies are applying SOX principles"

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[2] Due to the considerable obstacles in obtaining access to convened audit committees for experimentation purposes, we conduct our experiment at an individual versus group level. While we use the term "audit committees" to refer to audit committee members, we acknowledge that our findings relate to individual committee members, not audit committees per se.

[3] These participants likely are more representative of sophisticated versus unsophisticated investors; yet for parsimony, we refer to our participants as "investors."

(PricewaterhouseCoopers 2005). Another reported that 43 percent of private companies have or plan to adopt policies concerning the pre-approval of nonaudit services (Hartman 2004).

This renewed emphasis on an audit committee's responsibility for the audit rekindles a longstanding debate about the effects of the auditor's provision of nonaudit services on audit quality. On the negative side of the debate are those who believe joint service provision places the auditor in a compromised position (e.g., in one of evaluating his or her own work) or in a position of management (e.g., SEC 2000). In support, some academic studies find an association between nonaudit services and lower audit quality due to impaired auditor objectivity in trying to retain nonaudit service fees (Frankel et al. 2002; Ferguson et al. 2004). Regardless of the mechanism, the overall concern is that joint service provision can impair the auditor's objectivity and result in lower audit quality.[4]

On the other side of the debate are those who believe that joint provision of audit and certain nonaudit services improves audit quality. One supporting position is that some nonaudit services provide the auditor with essential knowledge of a client's business and transactions (Copeland 2002; Goldwasser 2002; Pletz 2002; Shedlarz 2002) and thus improve audit effectiveness via knowledge spillovers (Simunic 1984; Whisenant et al. 2003; Kinney et al. 2004; Flaming 2005). For example, risk management services promote an effective risk assessment, an analysis critical to evaluating audit risk and ensuring audit quality (Goldwasser 2002; Rice 2000). A second argument favoring joint provision is that it increases audit firm intellectual capital, which, in turn, improves the overall quality of the audit.[5] Anecdotal evidence suggests that many corporate boards do see such benefits as they support significant levels of nonaudit services (see, Weil and Tannenbaum 2001; Frankel et al. 2002). Regulators also recognize that some nonaudit services may in fact improve audit quality (Federal Register 2003). Thus, there is an acknowledged trade-off between the possible impairment of auditor objectivity and the potential benefits to audit quality from the auditor's provision of nonaudit services.

In pre-approving nonaudit services, the audit committee should consider the trade-off between auditor independence (objectivity) and audit quality and be prepared to justify its decision to those to whom it is accountable: the board, management, and private/public investors (Braiotta 2004). Research in both psychology and accounting indicates that this accountability and potential for justification may impact audit committee decisions.

Audit Committee Accountability

Accountability results when individuals have implicit or explicit expectations that they may be called upon to justify their beliefs, feelings, or decisions (Lerner and Tetlock 1999). Failures to provide satisfactory justification are assumed to be associated with negative consequences; however, the consequences need not be significant or tangible for accountability to occur (Siegel-Jacobs and Yates 1996; Lerner and Tetlock 1999). Individuals held accountable for their evaluation process typically consider more of the available information and think more deeply about the reasoning that led to their decisions. In addition, such individuals tend to engage in a more critical evaluation of decision choices and decision-related trade-offs (Tetlock et al. 1989) and a less biased evaluation of decision alternatives (Peecher and Kleinmuntz 1991; Lerner and Tetlock 1999).[6]

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[4] There is general agreement that if a nonaudit service places the auditor in a compromised position, then the auditor cannot remain objective. However, it is a personal judgment as to which specific nonaudit services this category comprises. Because actual objectivity is not observable, individuals form perceptions about objectivity from experiences and views of others. Thus, it follows that individuals would have different thresholds for what would or would not compromise actual objectivity. We expect experience to influence these personal beliefs.

[5] Practitioners also associate the ability to perform high-quality audits with the breadth of firm-level expertise of those outside the audit function who are brought in for consultation (Copeland 2002; Plitch 2002). A companion self-selection argument is that the offering of both audit and nonaudit services makes the audit firm more attractive to highly qualified and talented professionals (Copeland 2002).

Responses to accountability can be influenced by whether the preferences of the audience to which one is accountable are known or unknown (Seidenfeld 2001; Lerner and Tetlock 1999). In general, individuals "are well aware of the preferences of those to whom they are accountable" (Pennington and Schlenker 1999, 1068) and when preferences are known, individuals tend to "alter the outcome of their decisions to come closer to the outcome that would satisfy their audience," especially if the outcome could be disagreeable (Seidenfeld 2001, 1066). This is true even if the audience preferences are not explicitly known, but can be deduced (Tetlock et al. 1989; Seidenfeld 2001). In sum, knowing audience preferences leads individuals "to anticipate objections that others may raise to the positions they are about to take" (Tetlock and Boettger 1994, 1–2) and to select a course of action that best conforms to the audience's views (Tetlock 1983; Tetlock et al. 1989).

Given the audit committee's accountability to third parties for the oversight of the audit process and audit quality, it is reasonable that the committee engage in a cost-benefit analysis to determine whether to purchase nonaudit services from the auditor. That is, in making a recommendation to the board and management, a committee likely would consider the benefits of joint provision (e.g., improved audit quality, the desired outcome) relative to the costs (e.g., compromised audit quality due to auditor independence impairment). If an audit committee believes a particular auditor-provided nonaudit service is incompatible with auditor independence, then the potential costs to audit quality significantly outweigh the benefits. Thus, the committee would recommend against joint service (i.e., hire an unaffiliated firm). This position is justifiable in terms of audit quality so there could be little objection from third parties.

If, on the other hand, joint provision is seen as compatible with auditor independence, then the audit committee would be expected to evaluate the decision alternatives relative to the effect of the nonaudit service on audit quality. A committee should be willing to approve the purchase of nonaudit services from the auditor if the joint provision is believed to enhance audit quality. This position could be justified given higher audit quality is in the best interests of the company. Alternatively, if the nonaudit service does not improve audit quality, then the decision to hire the auditor is more difficult to justify, and audit committees are less likely to recommend joint provision.[7]

Assuming no auditor objectivity concerns, it seems reasonable for audit committees to recommend the purchase from the auditor of allowable nonaudit services that are believed to improve audit quality. However, presently no empirical evidence exists on whether audit committees actually consider the effects on audit quality in their nonaudit service approval decisions. Thus, we test the following hypothesis:

> **H1a:** *Ceteris paribus*, if audit committees believe joint provision improves audit quality, then they will be *more likely* to recommend the auditor for provision of nonaudit services.

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[6] We thank a legal scholar for suggesting that audit committees are clearly aware that they may be called upon to explain to judges the evaluation *process* itself should legal issues arise related to a decision.

[7] The reduction in overall costs if all services are provided by a single firm is one potential justification for approving joint provision. However, justification on this argument alone may be less defensible as there could be concerns of potential independence issues. Thus, a recommendation based solely on a cost-savings argument would be likely viewed as open to criticism.

Audit Committee Responsibility to Investors from Public Disclosures of Committee Decisions

Along with their responsibilities to the board, management, lenders, and other owners, public company audit committees have increased accountability to regulators and investors. Recent regulatory actions targeted at investors' perceptions demonstrate the importance of the audit committee's responsibility to instill investor trust. For example, as part of the governance reforms, the SEC passed two additional disclosure requirements aimed at influencing investor perceptions. The first requires companies to report the audit and nonaudit fees paid to the auditor (SEC 2000; Federal Register 2003).[8] The SEC states that such disclosures allow investors to decide whether the auditor is in a position to give an objective opinion on the financial statements (Unger 2001) and have "the capacity to shape the behavior of investors, possibly decreasing investors' confidence in the validity of their financial statements" (Rittenhouse 2001, 1).

A second requirement relates to the disclosure of the audit committee's pre-approval process and its consideration of whether the joint provision of services is compatible with the auditor remaining objective (Federal Register 2003). While the act of pre-approval holds audit committees directly accountable for actual auditor independence (and thus, audit quality), its disclosure is aimed at increasing investors' confidence in the auditor independence evaluation process (Raghunandan and Rama 2003; Federal Register 2003). The SEC contends that such "disclosures ... allow investors to ask more direct and useful questions of management and directors regarding their decisions to engage the accountants for such services" (Federal Register 2003, 6037).[9]

The act of disclosing the consideration of the compatibility of the nonaudit service with auditor independence underscores the fact that audit committees are accountable for the process they use in approving joint provision including an evaluation of auditor independence. But this accountability is now to the regulators who "believe that nonaudit services (NAS) may cause the auditor to be perceived as 'dependent' in appearance ... even if they have no impact on the quality of financial statements" (Nam 2005, abstract), and investors who will form opinions of audit quality based in great measure on their *perceptions* of auditor independence. Because audit committees are aware that the SEC intends for these disclosures to influence investors' perceptions of auditor independence, justification of the decision would include concerns related to *perceived* auditor independence as well as audit quality.

Based on prior research in accountability, we expect audit committees who are held publicly accountable to regulators and investors through disclosures to be more concerned with the ability to justify the decision from an auditor independence standpoint.[10] Because of the nature of the disclosures and of the press surrounding the politics underlying the mandates, we expect audit committees to view disclosures as having *negative* implications with respect to another member of the audience, investors, and to be more specifically concerned with investors' perceptions of auditor independence (see, Flaming 2005). Thus, audit committees are likely to view disclosures as having downside risk with little upside potential. Because the pre-approval process and nonaudit service fees are publicly disclosed only if the audit committee recommends the auditor for nonaudit services, they will be less likely to pre-approve the purchase of nonaudit services from the auditor. This is consistent with the SEC's contention that disclosures may raise concerns about independence impairment perceptions. Specifically, "[c]ompanies may perceive disclosing the ratio of fees for non-audit to audit services [to] decrease investor confidence ... [and in] that case, we can expect companies to take steps to improve that ratio" (Unger 2001, 3). This would be consistent with companies such as International Paper Co. having established "bright lines" in which the purchase of any nonaudit service that brings the ratio of nonaudit fees to audit fees over a specified threshold is rejected (Goldwasser 2002; Plitch 2002; Bryan-Low 2003).[11]

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[8] Section 10A(i)(2) of the Exchange Act entitled "Disclosure to Investors" requires the company to disclose the amount of fees billed for the two most recent years by the auditor in four categories: audit, audit-related, tax, and all other. In addition, the company must disclose the types of services provided under all categories other than audit services (Federal Register 2003).

[9] In this study, we did not try to separate out the pre-approval versus fee disclosure effects as they naturally coexist in practice, and actual corporate directors would have experienced these together as well. Overall, from SEC deliberations with which we are familiar, the expected effects of disclosures, particularly related to fees, are likely negative.

[10] This accountability would lead individuals to "avoid acting in ways that [will] trigger ... [blame] from well-defined constituencies" (Tetlock and Boettger 1994, 4) and to weight more heavily factors related to perceived independence (see, e.g., Peecher 1996).

Therefore, we propose the following hypothesis:

> **H1b:** *Ceteris paribus*, audit committees will be *less likely* to recommend the auditor for joint provision of nonaudit services when public disclosures are required relative to when disclosures are not required.

Possible Interactive Effects between Audit Quality and Disclosures

As previously discussed, we believe audit committees will engage in a cost-benefit analysis whereby the potential benefits to hiring the auditor are evaluated relative to the costs. If committees believe that joint provision impairs the auditor's objectivity, then the costs exceed any potential benefits to audit quality and an unaffiliated firm should be hired for the nonaudit services. Anecdotal evidence, however, shows that audit committees are willing to approve joint service provision, consistent with the belief that there are benefits (e.g., to audit quality or at least no significant costs) in the auditor providing some nonaudit services. If actual auditor independence is uncompromised, then we expect audit committees to approve joint provision more often when audit quality improves (H1a). In addition, the willingness of audit committees to recommend joint provision, all things equal, should be greater in the private company setting as there are no additional costs associated with disclosure (H1b).

In light of accountability to public parties, however, we expect audit committees to more critically and differentially consider the downside of public disclosures on regulators' and investors' perceptions of auditor independence relative to the benefits to audit quality. We base this on the increased accountability to well-defined and important constituencies and a basic cognitive tendency for "losses to loom larger than gains" (Tetlock and Boettger 1994, 4). An overriding concern with disclosures regardless of potential audit quality improvements would be consistent with the observation that some corporate boards (e.g., Johnson & Johnson) take the approach of simply refusing to purchase any additional services from their auditor (Goldwasser 2002; Bryan-Low 2003).

Thus, given the disclosure downside, we expect audit committees' willingness to recommend joint provision when audit quality improves to diminish. Therefore, we examine the extent to which audit committees are willing to forgo audit quality to avoid disclosure and propose the following hypothesis:

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[11] We have no knowledge of any specified relative "bright line" amounts or relations; however, a few companies have indicated that nonaudit service fees should not exceed audit fees.

H1c: Audit committees are *more likely* to recommend joint provision that improves audit quality relative to when joint provision does not improve audit quality. However, the likelihood of recommending joint provision due to audit quality improvement is moderated by public disclosure.

III. METHOD

Design

To address these three research questions (H1a–H1c), we employ a 2 × 2 between-participants design. The first factor relates to whether the proposed nonaudit service improves audit quality. This factor is operationalized by varying the type of nonaudit service: If provided by the auditor, then risk management services are expected to result in a higher audit quality. In contrast, there is no likely audit quality effect for the joint provision of human resource services.[12] The second factor is whether the company is required to publicly report the mandated disclosures associated with joint provision of audit and nonaudit services. This presence or absence of a disclosure requirement is operationalized by the type of company, public or private, respectively.

Participants

One hundred corporate directors who were attending one of 12 different corporate governance roundtables held around the United States participated in the study. On average, these directors are serving or have served on 3.2 corporate boards and have 5.3 years (3.2 years) experience as a public (private) company audit committee member.[13] Summary background information is presented in Table 1.

Procedure

We worked with a national coordinator of the governance seminars as well as the individual administrators at each of the 12 locations to distribute and collect the instruments.[14] Each location administered instruments for all experimental conditions; thus, participants were randomly assigned to conditions at each location. Upon arrival at a seminar, participants received a packet comprising: (1) a short letter of introduction asking for their voluntary participation in the study that would address an audit committee's approval for the client's purchase of nonaudit services; (2) two pages of background information on the

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[12] We specifically selected risk management (human resource) services as they are representative of services that typically are viewed as improving (having little effect on) audit quality. We based our selection on reviews of auditing firms' publications, statements from the NACD website, and conversations with affiliates from KPMG's Audit Committee Institute. Even though these services' effects on audit quality are generally viewed this way, and we instructed participants that they could be confident that risk management (human resource) services would improve (have no effect on) audit quality, it is possible for an individual's experiences to alter his or her views. To that end, we collected data to assess the extent to which participants' views matched their assigned conditions.

[13] A total of 103 corporate directors participated; however, we dropped from the study two non-director (academic) participants and one director who failed the manipulation check question. As subsequently discussed, 81 of these 100 directors participated as audit committee members while 19 directors acted in the capacity of investors. We collected audit committee experience data on small-, mid-, and large-cap firms and also gathered the number of years of audit experience. The only significant difference in experience between groups was that 31.6 percent of the 19 investor participants had more than ten years of audit experience compared to the 20.0 percent of the audit committee participants in the public-risk audit committee condition. Inclusion of overall or individual experience covariates does not alter any of the reported results or interpretations.

[14] The national coordinator assisted in drafting a uniform set of instructions to all administrators and attended six of the seminars to ensure consistency in the distribution and request for participation. When the national coordinator was not in attendance, he assisted in the writing of the administrators' announcement such that identical statements would be made in all locations.

TABLE 1
Sample Size and Demographic Data for Corporate Director Participants

Panel A: Sample Size

Total number of corporate director participants	103
Less: Participants with an academic versus director background	(2)
Participant who failed the manipulation check	(1)
Final sample	100

Panel B: Corporate Director and Audit Committee Experience Data for All Participants

	All Participants (n = 100)	Audit Committee Members (n = 81)	Investors (n = 19)
Average years on private company audit committees	3.2	3.3	3.2
Average years on public company audit committees	5.3	5.4	4.9
Average number of directorships in other companies	3.2	3.0	3.7

company and decision; (3) a one-page question set related to the primary decision; and (4) a two-page question set containing manipulation check, explanatory, and demographic questions. Participants were instructed to return the completed instrument to the administrator at the end of the seminar.[15]

Task and Instrument

Each of the 81 participants was instructed to assume the role of an audit committee member for a growing small-cap company. They were told that at its upcoming meeting, the audit committee would be recommending one of two Big 4 firms to provide some nonaudit services to the company and that they would be one of two audit committee members to whom the initial selection decision has been delegated.

All participants learned that Firm A (the external auditor) was just hired to conduct the company's external audit and would be paid audit fees of $1.3 million. They were also told that Firm A "has a well-established reputation for providing these [nonaudit] services to some of the leading firms in [the industry]" and has "access to relevant best practices." In contrast, participants were told that the second firm, Firm B, had no present or prior association with the company, but did have experience in providing these nonaudit services. Both accounting firms bid $2 million for the 18-month nonaudit service project, with $1.25 million expected to be billed this year.[16] We wanted Firm A to appear slightly more attractive in terms of the quality of the services (or at least no less attractive) and equivalent in costs such that selecting Firm B would be for reasons other than service quality or cost savings.[17]

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[15] The materials included a postage-paid, addressed, return envelope for participants to return the instrument via mail, if necessary. We received fewer than 15 percent of the 103 instruments in this fashion with no concentration in any one condition.

[16] We reviewed audit fee and nonaudit fee disclosures for a sample of 41 small-cap companies to determine a reasonable amount for the nonaudit service fees as well as the relative amount of the nonaudit to audit fees. We also had two partners from different Big 4 firms review the instruments for reasonableness and realism. We used an approximate 1:1 ratio of audit to nonaudit service fees; this is considerably smaller than that reported by Frankel et al. (2002) and Weil and Tannenbaum (2001) and used by Flaming (2005).

The instrument for each experimental condition comprised two pages of background information. Page one described the nature of the *allowable* nonaudit service and the relation of the nonaudit services to audit quality. In the risk management services cases, we told participants they could be confident that a higher quality audit would result if Firm A provided both services. Participants in the human resource service cases were told that purchasing joint services from Firm A would not lead to a higher quality audit.

Page two of the background information varied depending on whether the company was publicly or privately held. For those assigned to the public company (traded on NASDAQ) condition, the second page presented alternative sets of public disclosures that would be reported in the company's filing based upon the committee's recommendation: those required if the audit committee recommended the audit firm (Firm A) for the nonaudit services and those required if the unaffiliated firm (Firm B) provided the nonaudit services.[18] In both cases, disclosures would show that \$1.3 million was paid in audit fees. If the audit firm was selected to provide the nonaudit services, then the company would report \$1.25 million in other fees paid to the auditor. In addition, disclosures would state that the audit committee pre-approved all services purchased from the auditor. Alternatively, if the unaffiliated firm (Firm B) was selected to provide nonaudit services, then the company would report \$0 in other fees paid to the auditor and there would be no mention of the pre-approval process (or even the fact that the company purchased any nonaudit services).

For those receiving the private company condition, page two of the background information stated the total amount that the company would pay for audit and nonaudit services, emphasizing that the amount would not differ; the only difference is whether the company pays \$1.25 million for nonaudit services to Firm A or Firm B. In addition, the instrument indicated that because the company was privately held, it would not be required to make any public disclosures regarding the amounts paid for audit or nonaudit service fees nor would it be required to disclose that the audit committee pre-approved such services.

After reading the background information, participants answered two sets of questions. The first question set asked participants to indicate which firm they would recommend to provide the nonaudit service. They were then instructed that they might have to explain the reasoning behind their recommendation and, thus, were asked to list: (1) the single most important factor in support of the recommended firm, (2) other factors they considered that support their recommendation, and (3) the single most important factor that argues against the recommended firm. After inserting this question set into an envelope, participants responded to a second set of manipulation check, control, and demographic questions.[19]

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[17] Based on pilot results and debriefing sessions, we believed it necessary to control for these factors in order to eliminate these reasons as alternative explanations.

[18] The SEC requires disclosure of amounts paid to the auditor for audit fees and for various categories of other services, e.g., audit-related, IT, and tax services. In addition, if the company purchases nonaudit services from the auditor, additional disclosure that the audit committee pre-approved such services is required. The disclosures presented in the instrument were patterned from several actual small-cap firms' 10-K filings that disclosed nonaudit services. We presented the disclosed audit and nonaudit fees in two categories: audit fees and "all other fees."

[19] Twelve accounting faculty, several accounting Ph.D. students, and 60 M.B.A.s (with an average of 62 months of work experience) pilot-tested the instrument for task and questions clarity, completeness of information, and timing. Two former audit partners and the national coordinator of the seminars reviewed the revised instrument for realism, reasonableness, and length. We made a few additional revisions based on these comments.

IV. AUDIT COMMITTEE RESULTS

Manipulation Checks, Control, and Test Variables

Participants responded either "yes" or "no" to a manipulation check question that asked whether the case materials *indicated* that the joint provision of the nonaudit service and audit would result in a higher quality audit. Those who received the risk management service cases should have responded "yes," whereas participants receiving the human resource service cases should have answered "no." One participant responded incorrectly and, accordingly, was eliminated from all analyses.[20]

The ability to examine whether audit committees consider the effect on audit quality when recommending the auditor provide nonaudit services depends on committees *believing* that the joint service provision does or does not improve audit quality. Given the backgrounds of the majority of our participants, it is likely that their prior director experiences and/or real-world prior opinions could dominate what they were instructed regarding audit quality and could influence the extent to which they *believe* audit quality improves with any given nonaudit service. Thus, we asked participants "to what extent do you *believe* the auditor's provision of the risk management (human resource) services would affect audit quality (AQ)?"

Participants responded on an 11-point scale with 1 being "AQ would be much lower," 6 (the mid-point) being "no effect on AQ," and 11 being "AQ would be much higher." Those assigned to the human resource conditions ($n = 40$) reported a mean (standard deviation) of 6.13 (1.18). This is not significantly different from the "no effect" of 6.0 ($p = 0.51$). In comparison, mean response for participants in the risk management conditions ($n = 41$) was 7.32 (1.84), which is significantly greater than the "no effect" of 6.0 ($p < 0.001$). While the two means also are significantly different ($p < 0.001$), the distributions for these responses suggest that some participants believed that risk management (human resource) services had no measurable effect on (improved) audit quality. Thus, in our analyses, we use a measured quality belief variable (i.e., whether participants believed joint provision did or did not improve audit quality) as the independent variable instead of the manipulated service condition.[21]

Using a measured variable rather than the manipulated variables as the independent variable limits one of the advantages of experimentation, randomization. However, Libby et al. (2002, 24–25) support the use of measured variables in at least four circumstances, several of which are relevant in our setting. First, it is not always possible to manipulate an antecedent; second, the theory involves a mediating belief variable, one that cannot be manipulated; third, it is often less interesting to examine reactions to manipulated versus naturally occurring variables; and fourth, the measured variable provides keys to understanding an underlying process. In our setting, we are interested in examining beliefs—and how audit committees perceive the relation between a nonaudit service and audit quality may be influenced by prior experiences (which we cannot directly measure or control). By using the manipulated variables, we would add noise to our tests and reduce the chance of detecting the effects on latter judgments (Peecher and Solomon 2001; Libby et al. 2002).

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[20] We did not elicit a disclosure manipulation check. First, there were specific references and questions directed at the assigned public or private condition throughout the questionnaire. Second, pilot tests revealed no confusion regarding this construct. Finally, participants' explanations for their firm recommendations included disclosure-related reasons for the public company; no one in the private condition indicated disclosure as a basis for their recommendation.

[21] Nine of the 40 (14 of the 41) participants assigned to the human resource (risk management) condition responded differently than the intended service/audit quality manipulation.

This said, we acknowledge that using measured versus manipulated variables limits our ability to interpret any observed relations as causal—i.e., we can only discuss associations.[22]

As part of their pre-approval decisions, audit committees must assess whether joint service provision is compatible with auditor independence. If a service is viewed as incompatible, then the committee should not approve the joint provision. Thus, we asked participants to indicate how the joint provision of risk management (human resource) services would affect the auditor's actual objectivity on an 11-point scale with 1 being "Objectivity not affected," 6 (the mid-point) being "Objectivity slightly reduced," and 11 as "Objectivity significantly reduced."

Across all participants, the mean response (standard deviation) was 4.96 (2.87), indicating that, on average, participants view the joint provision of nonaudit services as slightly reducing auditor objectivity; the means for risk management and human resource services conditions were 5.05 (2.92) and 4.86 (2.85), respectively. In addition, those in the public and private company setting reported means of 5.12 (2.94) and 4.75 (2.81), respectively. Differences across service type and company type were not significant ($p = 0.77$ and $p = 0.60$, respectively)

Because we test the hypotheses using the measured versus manipulated variable, we also report the means for objectivity based on the quality belief variable. The means for those who "do not believe" and "believe" the provision of nonaudit services affects audit quality were 5.32 (3.02) and 4.50 (2.63), respectively. Differences across quality belief and company type were not significant ($p = 0.20$ and $p = 0.56$, respectively).

Finally, we attempted to eliminate nonaudit service quality as a possible explanation for the selection of Firm B by making the auditor (Firm A) a slightly more attractive nonaudit service provider. However, as with audit quality, nonaudit service quality beliefs may vary across individuals and possibly across types of services. As such, we asked participants which firm they believed would provide a higher quality nonaudit service or whether the firms were essentially equivalent in this regard. Thirteen participants indicated that Firm B would provide a higher quality nonaudit service; there were no significant differences across conditions or services between frequencies for those responding in favor of Firm B.

Descriptive Data for H1a, H1b, and H1c

Table 2 reports the descriptive data used to test H1a–H1c. Specifically, Panel A presents firm recommendations by type of company and audit quality beliefs for the 81 participants.[23] Overall, the descriptive data show that audit committees appear reluctant to hire the auditor for nonaudit services (only 37.0 percent of the time) in spite of the auditor having a slight advantage in providing the nonaudit services. However, if audit committees believe joint provision benefits audit quality, they recommend the auditor 52.8 percent of the time compared to only 24.4 percent of the committees when the service results in no audit quality improvements. Thus, while audit committees, on average, prefer joint service provision less than 40 percent of the time, they are more willing to hire the auditor when there are benefits

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[22] Given the calls for research on these issues, we believe this trade-off is justified. That is, the benefits obtained by the use of actual audit committee members (versus students or auditors acting as audit committee members) would seem to outweigh the limitations at this stage. Presently, archival data cannot address these research questions, and there is a paucity of research using actual audit committee members.

[23] We show two conditions for these beliefs: audit quality does not improve (all those reporting values less than or equal to 6.0) and audit quality improves (all values greater than or equal to 7.0). Relatively few values fall outside the range of 4.0 to 9.0; four of the 45 participants in the "does not improve" condition reported values less than 4.0 and only three of the 36 participants in the "improves" condition reported values greater than 10.0.

TABLE 2
Percentage of Audit Committee Participants Who Recommend the External Auditor (Firm A) Provide the Nonaudit Service

Panel A: Audit Committee Participants' Firm Recommendations by Audit Quality Belief and Type of Company

	The Effect of Nonaudit Service on Audit Quality Belief[a]		
Type of Company	**Service Does Not Improve Audit Quality**	**Service Improves Audit Quality**	**Overall**
Private (no disclosure)	28.0% (7 out of 25)	72.7% (8 out of 11)	41.7% (15 out of 36)
Public (required disclosures)	20.0% (4 out of 20)	44.0% (11 out of 25)	33.3% (15 out of 45)
Overall	24.4% (11 out of 45)	52.8% (19 out of 36)	37.0% (30 out of 81)

Panel B: Audit Committee Participants' Firm Recommendations by Type of Nonaudit Service and Type of Company

	Type of Nonaudit Service		
Type of Company	**Human Resource (no effect on audit quality)**	**Risk Management (improves audit quality)**	**Overall**
Private (no disclosure)	36.8% (7 out of 19)	47.1% (8 out of 17)	41.7% (15 out of 36)
Public (required disclosures)	23.8% (5 out of 21)	41.7% (10 out of 24)	33.3% (15 out of 45)
Overall	30.0% (12 out of 40)	43.9% (18 out of 41)	37.0% (30 out of 81)

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[a] Audit Quality Belief was measured on an 11-point scale. Those reported under "Service Does Not Improve Audit Quality" indicated values for Audit Quality Belief less than or equal to the mid-point of 6.0 (no effect). Those reported under "Service Improves Audit Quality," are those who reported values greater than 6.0.

to audit quality. This is consistent with H1a and the Public Oversight Board (POB) Panel on Audit Effectiveness's recommendation that audit committees consider audit quality in evaluating auditor independence (POB 2000).

Panel A of Table 2 further indicates that 41.7 percent of audit committees recommended joint provision when the company is private, relative to only 33.3 percent when public disclosures are required; this is consistent with H1b. This pattern holds within each audit quality belief condition. However, the increase in participants' recommendations of joint provision when audit quality improves versus does not improve is greater for the private company. When the company is private, the willingness of participants to recommend joint provision increases from 28.0 percent to 72.7 percent; in comparison, when the company is public, recommendations for joint provision increase from 20.0 percent to 44.0 percent when participants believe that the services will result in improved audit quality. These data are consistent with disclosure moderating the effects of audit quality beliefs (H1c).[24]

Tests of H1a, H1b, and H1c

We use the logistical regression model below to test the relation between nonaudit service, audit quality, and public disclosures on audit committees' firm recommendations (*FIRMREC*):

$$\begin{aligned}FIRMREC = {} & b_0 + b_1 AQBELIEF + b_2 PUBLIC + b_3 PUBLIC \times AQBELIEF \\ & + b_4 OBJBELIEF + b_5 NASQBELIEF + \varepsilon\end{aligned}$$

The dependent variable, *FIRMREC,* equals 1 if the audit committee member selects the audit firm to provide the nonaudit service, 0 otherwise. The main and interactive effects of the two independent variables are modeled, each measured as an indicator variable: *AQBELIEF* captures whether the participant believes the joint provision of the nonaudit service will or will not improve audit quality, and *PUBLIC* captures whether disclosure of jointly provided nonaudit services is or is not required.

In addition to our test variables, we control for the effects of two other factors that may affect participants' firm recommendations: a variable that captures participants' beliefs of the effect of joint provision on auditor's objectivity (*OBJBELIEF*) measured on a 1 (Objectivity not affected) to 11 (Objectivity significantly reduced) scale, and (2) an indicator variable (*NASQBELIEF*) that captures whether participants believe Firm B would provide higher nonaudit service quality.

The coefficient associated with the main effect of *AQBELIEF* (b_1) measures the effect of audit quality beliefs on firm recommendations for those participants whose companies are private, i.e., not required to make public disclosures about joint service provision. This same relation for participants whose companies are required to make such public disclosures is represented by the sum of the coefficients for this main effect variable and the interaction variable ($b_1 + b_3$). Given we posit that committees are more (less) likely to recommend the auditor when joint provision is believed to improve (not improve) audit quality, we predict positive coefficients for b_1 and the sum of $b_1 + b_3$. However, disclosures are predicted to moderate participants' willingness to recommend joint provision that is due to improved audit quality. Thus, we also expect a negative coefficient on the interaction between *AQBELIEF* and *PUBLIC* (b_3).

The coefficient for *PUBLIC* (b_2) measures the effect of disclosures on participants' firm recommendations for those participants who believe audit quality does not improve. This same relation for participants who believe joint provision improves audit quality is represented by the sum of the coefficients for this main effect variable and the interaction ($b_2 + b_3$). We expect negative coefficients for b_2 and ($b_2 + b_3$): audit committees will be less likely to recommend the auditor when disclosures are required.

Finally, we expect negative coefficients for *OBJBELIEF* (b_4) and *NASQBELIEF* (b_5): audit committees will be less likely to recommend the auditor when they believe joint provision compromises the auditor's objectivity or when they believe Firm B will provide a higher quality nonaudit service.

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[24] As a comparison, we present firm recommendations by type of service and company in Panel B of Table 2. While the predicted main effects are directionally consistent with that reported in Panel A, the differences are not as significant. Interestingly, the comparison between panels does suggest that audit committees are reacting to the influence of the service on audit quality and not service type itself.

As shown in Table 3, the overall Logit model is significant ($p < 0.001$) with a pseudo-R^2 of 42 percent. As expected, for the control variable *OBJBELIEF*, we see that participants are significantly less likely to recommend Firm A when they believe joint provision adversely affects the auditor's objectivity ($p < 0.01$). This finding is consistent with the SEC's mandate that audit committees consider the compatibility of nonaudit services with the auditor's independence. We do not find a significant relation between the participants' beliefs about nonaudit service quality and their firm recommendations. This may be due to the fact that the nonaudit service quality for those believing quality was higher for Firm B was a by-product of participants' beliefs about joint provision compromising auditor objectivity.[25]

After controlling for participants' beliefs that joint service provision may compromise auditors' objectivity, we find that participants are more likely to recommend Firm A when they believe joint provision will improve audit quality. The effect for *AQBELIEF* is significant when public disclosures are not required ($b_1 > 0$; $p = 0.02$) as well as when public disclosures are required ($b_1 + b_3 > 0$; $p = 0.05$). Together, these results support H1a and the POB's recommendation that audit committees consider the effects of nonaudit services on audit quality in their pre-approval decisions.

Hypothesis 1b predicts that participants will be less likely to recommend the auditor when public disclosures are required; it is only partially supported as the effect of disclosures on firm recommendations is mixed depending on whether joint provision improves audit quality. Specifically, disclosures do not have a significant effect on recommendations for those participants who believe joint provision does not improve audit quality ($b_2 < 0$; $p = 0.90$). While only 20.0 percent of participants are willing to recommend joint provision when disclosures are required, relatively few participants are willing to recommend joint provision even absent disclosures when there are no benefits to audit quality. The slight advantage of hiring the auditor in this case, i.e., the quality of the *nonaudit* service, is insufficient justification for many to recommend joint provision. In contrast, participants who believe audit quality improves are significantly less likely to recommend joint provision when public disclosures are required versus not required ($b_2 + b_3 < 0$; $p = 0.05$). This result is consistent with participants being willing to forgo potential benefits to audit quality to avoid the negative effect of disclosures on investors' perceptions.[26]

Hypothesis 1c predicts that disclosure moderates participants' increased willingness to recommend joint provision because of the benefits to audit quality. Overall, we expect participants to be more willing to recommend joint provision if audit quality improves; however, we expect less effect from improved quality when the joint provision will be

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[25] As an example, several participants viewed joint provision as impairing independence and, thus, believed this would reduce the quality of both the audit and nonaudit services. Others believed nonaudit service quality for Firm B would be higher just because there were "two sets of eyes." If participants recommended Firm B for the nonaudit services, then we intended it to be for reasons other than the quality of its services. It appears from this result that it may indeed be the case. We did, however, also analyze our model of firm recommendations (excluding the *NASQBELIEF* variable) for just those participants who believed Firm A provided equivalent or higher quality nonaudit service. For this subset of $n = 68$, the coefficient estimates and their significance were identical to the full sample.

[26] Thirty-nine of the 45 participants assigned to the public condition responded "yes" that required nonaudit service disclosures would affect investors' perceptions of auditor independence. The mean response (standard deviation) on an 11-point scale that investors would perceive the auditor as "slightly less independent" ("1") or "not independent," ("11") for these 39 participants was 6.05 (2.56); this indicates that, on average, audit committees believe disclosures negatively influence investors' perceptions of auditor independence as purported by the SEC. While there was a difference in means between nonaudit service types, mean = 7.31 and 5.17 for human resources and risk management, respectively, the difference was not significant between audit quality belief conditions (mean = 6.88 and 5.48 for audit quality not improved and improved, respectively).

TABLE 3

Logistic Regression of Audit Firm Recommendation on Audit Quality Belief, Type of Company, Auditor Objectivity Belief, and Nonaudit Service Quality Belief

Tests of H1a and H1b

$$FIRMREC = b_0 + b_1 AQBELIEF + b_2 PUBLIC + b_3 PUBLIC \times AQBELIEF + b_4 OBJBELIEF + b_5 NASQBELIEF + \varepsilon$$

Variable	Predicted Relation	Estimated Coefficients	Wald Chi-Square
Intercept	none	1.712	4.419**
AQBELIEF	+	2.529	4.178**
PUBLIC	–	−0.842	0.898
PUBLIC × *AQBELIEF*	–	−1.098	0.535
Joint Test ($b_1 + b_3$)	+	1.431	2.736**
Joint Test ($b_2 + b_3$)	–	−1.940	2.556**
OBJBELIEF	–	−0.494	14.465***
NASQBELIEF	–	−14.959	0.003
Number of Observations	81		
Chi-Square for Model (5 degrees of freedom)	44.476		
p-value	0.0001		
Pseudo R^2	0.42		
Concordant Pairs	89.3%		

, * Statistically significant at less than the 0.05 and 0.01 levels, respectively, based on one-tailed (two-tailed) tests for variables whose relation to the dependent variable is (is not) predicted.

FIRMREC = 1 if participant recommends Firm A provide the nonaudit service, else 0;
AQBELIEF = 1 if participant believes the joint provision of the nonaudit service would improve audit quality, else 0;
PUBLIC = 1 if participant is assigned to the public company condition (disclosures required), else 0;
OBJBELIEF = extent to which the participant believes the joint provision of the nonaudit service would affect the auditor's actual objectivity, based on an 11-point scale (1 = objectivity not affected, 11 = objectivity significantly reduced); and
NASQBELIEF = 1 if participant believes Firm B will provide a higher quality nonaudit service, else 0.

disclosed. While the frequencies reported in Table 2 and the significance levels between b_1 and ($b_1 + b_3$) are directionally consistent with our predictions, the coefficient for the interaction between *AQBELIEF* and *PUBLIC* (b_3) is not significant ($p = 0.54$).

Additional Exploratory Analyses

Participants were told that it was "quite likely that another audit committee member will ask you to explain the reasoning behind your recommendation" and were asked to "list the *single* most important factor in support of the firm that you selected." While this question was open-ended, surprisingly, the breadth of responses was relatively low.[27] For the 24.4 percent of those participants who selected Firm A when joint provision did not improve audit quality, the most frequently indicated reason was that of the nonaudit service and/or provider quality (81.8 percent). The other reason provided was the synergies related to joint provision.[28] In comparison, there were three responses for the 52.8 percent of participants who recommended Firm A when audit quality improved: 42.1 percent indicated nonaudit service/provider quality; 36.8 percent indicated improved audit quality, and 21.1 percent recommended Firm A because of synergies.

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[27] Two individuals unaware of the research hypotheses coded the responses based on a provided template. The primary questions in coding related to whether "independence" referred to perceived or actual. In some cases, a determination could be made based on other responses. Because of the ambiguity in those remaining responses, we discuss independence concerns together.

Responses between participants in the private versus public company conditions differed only to the extent that those in the public-audit, quality-improved condition listed audit quality as their primary reason more frequently than those in the private-audit, quality-improved condition (46 percent compared to 25 percent, respectively). While the sample sizes are relatively small, this result seems to suggest that participants in the public setting were less willing than those in the private setting to recommend joint provision unless there were clear benefits to compensate for the potential negative effects of disclosures.

For participants selecting Firm B, the most commonly stated reason was a concern regarding independence issues. It was not possible in each case to determine whether the participant was concerned about actual or perceived independence; however, of those 51 participants recommending Firm B, 56.9 percent listed independence as their primary underlying reason for their selection. In addition to those who expressly mentioned independence, the remaining explanations also appeared independence-related: two indicated "two-firm benefit," two more listed "competition," and the remaining four mentioned nonaudit service quality and the audit/nonaudit service fee relation.

Finally, we asked participants whether changing the relative amount of audit to nonaudit service fees would change their recommendation. Approximately 57 percent of those who selected Firm A—and were not troubled with joint provision when the ratio was about 1:1—would change to Firm B if nonaudit service fees grew relative to audit fees. However, the findings also suggest that participants are not simply "running away from the audit firm" as 41 percent of those who selected Firm B would change to Firm A if nonaudit service fees were less.

V. INVESTORS' PREFERENCES

Hypothesis Development

In adopting the nonaudit service disclosure rules, the SEC stated as motivation the need to provide greater transparency to investors (Turner 2001; Pitt 2002). Specifically, the SEC's original proposal defended the fee disclosures as a way of providing "sunlight to the auditor independence area" (SEC 2000, Section 1) so investors can "determine for themselves whether there are concerns related to the auditor's independence" (Federal Register 2003, 6037). This disclosure policy seems consistent with research in accounting that demonstrates investors' decisions are influenced by information transparency (e.g., Hirst and Hopkins 1998; Maines and McDaniel 2000). In addition, the SEC believes investors may ask "more direct and useful questions" due to the increased transparency that disclosures bring (Federal Register 2003). Thus, based on prior accountability arguments, the direct link between audit committees' pre-approval decisions and required disclosures, and prior research, we believe audit committees will view this disclosure transparency as negative; by approving an unaffiliated firm for nonaudit services, audit committees can avoid this investor scrutiny. In addition, committees may reduce the possibility of regulatory scrutiny as the absence of nonaudit service disclosures would signal no auditor-provided-service conflicts of interest.

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[28] In most cases, this was not explained in any more detail. However, because the nonaudit service fee bid was equal across firms and audit quality did not improve, we assumed this related to nonaudit quality enhancing synergies, such as less audit committee time required for discussions with the firms.

While providing "sunlight," the fee disclosures are likely to have a negative, or at best, neutral effect on investors' perceptions of auditor independence (e.g., see, Dopuch et al. 2003; Flaming 2005). This effect may be due in part to the exact nature of these public disclosures. Specifically, investors learn of the amounts paid to the auditor for nonaudit services as compared to audit fees; while they also see that the audit committee approved such services, disclosures generally provide no indication as to why the auditor was hired. For example, was the audit firm the most qualified and/or the cheapest provider of the service, or were there synergies from hiring the auditor, e.g., improved audit quality? In other words, disclosures do not reveal all of the private information available to the audit committee.[29]

Investors rely on audit committees to use this private information to make decisions, and we assume investors prefer audit committees take actions that result in the highest audit (and financial reporting) quality. Yet, in spite of private information, if audit committees accept lower audit quality in order to avoid disclosures, an interesting question is whether audit committees really are protecting investors' interests. Unlike audit committees, investors face no public accountability with respect to fee disclosures; as a result, investors likely view joint service provision that improves audit quality as more positive relative to audit committees. That is, if investors receive the same "full" information available to audit committees, we would expect them to prefer joint provision more often than committees would be willing to recommend. We investigate whether audit committees' recommendations are consistent with investors' preferences and propose the following hypothesis:

> **H2:** *Ceteris paribus*, when joint provision improves audit quality and disclosures are required, compared to investors' preferences, audit committees will be *less willing* to recommend the auditors' provision of nonaudit services.

Participants and Method

To examine H2, we created a fifth experimental condition whereby 19 other director participants assumed the role of an investor in a public company. On average, these participants served on approximately 3.7 boards and for 3.2 (4.9) years on audit committees of private (public) companies. Summary background information is presented in Panel B of Table 1 and suggests that these individuals will be more similar to sophisticated investors.

These participants received packets comparable to the public company, risk management services, audit committee condition.[30] The audit firms and the risk management services information remained identical to that in the public-audit committee condition. After reading the background information, participants indicated the firm (Firm A or B) that they would prefer the audit committee recommend provide the nonaudit services; they also listed reasons for and against their selections. After inserting this question set into an envelope, participants proceeded to a second set of manipulation check, clarification, control, and demographic questions. These questions were nearly identical to those assigned to the audit committee role, albeit modified for the investor role.

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[29] Companies can voluntarily disclose the reason for hiring the audit firm; however, this practice appears to be a relatively rare occurrence.

[30] Because we were interested in *investors'* perceptions, there was no private company condition. Based on sample size considerations, we limited our examination to the case in which joint provision is expected to improve audit quality, i.e., risk management services.

Results

We asked participants how difficult it was for them to respond as investors rather than as audit committee members. The mean response (standard deviation) was 3.68 (2.95) on an 11-point scale with 1 and 11 being "not at all" and "extremely" difficult, respectively, suggesting participants had little difficulty responding as investors.

Second, participants responded to the question, "To what extent do you believe the auditor's provision of the risk management services would affect audit quality (AQ)?" On an 11-point scale (1 is "AQ would be much lower" and 11 is "AQ would be much higher"), these participants reported a mean of 6.68 (1.92), which was different from "no effect" of 6.0 at $p = 0.07$, suggesting investor participants believe audit quality improved slightly with the auditor's joint provision of risk management services.

Underlying the SEC's mandate is the belief that the disclosures allow investors to evaluate auditor independence. Parallel to the audit committee question, we ask to what extent the joint provision of services affects investors' perceptions of auditors' objectivity. Investor participants reported a mean (standard deviation) of 4.26 (2.47) suggesting that investors believe auditor objectivity is slightly reduced by the joint provision, consistent with the SEC's contention. This finding further suggests that audit committees' concerns over investors' perceptions of auditor objectivity may not be unfounded.

On average, 80.0 percent of investors prefer that the auditor provide risk management services that improve audit quality in contrast to audit committees' recommendations of only 44.0 percent of the time. This difference is significant at $p < 0.03$, consistent with H2. Interestingly, investors' preferences for joint service provision exceed audit committees' recommendations even though committees believe joint provision improves audit quality more relative to investors (mean *AQBELIEF* for investors is 6.68 compared to 7.83 for audit committees in the public-risk condition).[31]

For the investor participants who selected Firm A, the most frequently listed reason for their recommendations was nonaudit service/provider quality (62.5 percent). The other investors (37.5 percent) indicated improved audit quality as the basis for recommending the auditor. In addition, the majority of the investors who listed nonaudit service quality as their primary reason indicated audit quality as the second most important reason for selecting Firm A. In comparison, only two investor participants recommended Firm B; both indicated independence concerns as the reason for selecting an unaffiliated firm.

VI. DISCUSSION

One problem repeatedly cited in regulatory debates over rules restricting auditor provision of nonaudit services is the lack of empirical evidence indicating that joint provision, in fact, leads to actual and/or perceived auditor independence impairments (e.g., POB 2000; Turner 2001). Also questioned is the effectiveness of SEC-required disclosures in meeting the objectives of increased financial reporting quality. Former SEC Chairman Pitt (2002), in discussing the then-impending adoption of the Sarbanes-Oxley Act, expressed concern about the lack of empirical evidence supporting the new disclosure requirements. Further, he stated that if and when problems are empirically shown to exist, reforms should be tailored to fit those specific problems (Pitt 2002).

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[31] In this setting, audit committees made recommendations based on current disclosure rules; i.e., they know that investors have no additional information on audit quality, per se. Thus, we do not know whether audit committees' recommendations would change to be more consistent with those of investors if audit-quality-related disclosures were also required. We made this design choice knowing that access to participants was limited.

There have been other requests for research on policy implications. For example, the POB (2000) called for the continuous evaluation of the effectiveness of new SEC disclosure requirements as well as future rule-making initiatives. Former Acting SEC Commissioner Unger (2001) stated that the nonaudit service disclosure requirements "lay the groundwork for future study of the effect of nonaudit services on auditor independence" and expressed a need for "better empirical information on the effect of various nonaudit services in the future." Karmel (2005, 79) states, "[w]hether Sarbanes-Oxley will result in better corporate governance and greater sensitivity by corporate officers and directors to investor interests remains to be seen."

This research addresses such calls for evidence. Our results indicate that audit committees are reluctant to recommend the auditor for allowable nonaudit services if there are no direct benefits to the audit. This finding supports the contention that "audit committees are taking the pre-approval process very seriously ... and are appropriately restricting pre-approving services based on what is in the investors' bests interests" (Turley 2004, 3). The result holds even when audit committees have no public accountability, consistent with boards "of many private companies [embracing] various aspects of the [Sarbanes-Oxley Act]" (Gehrke 2006, 2) and the PCAOB's desires that standards can be adopted in private as well as public settings (McDonough 2004).

We further find that public disclosures not only seem to increase audit committees' sensitivity to *perceptions* of independence, but also enhance committee members' sensitivity to *actual* auditor independence issues. This is evidenced by the significant coefficient on our *OBJBELIEF* variable in our model, a result that is consistent with the SEC wanting greater board accountability for auditor independence. While audit committee members recommend an unaffiliated firm for nonaudit services when they view joint provision as incompatible with auditor objectivity (regardless of whether disclosures are required), others consider the possible negative effects of mandated disclosures on investors' perceptions. This finding is consistent with empirical research that demonstrates decisions become more conservative in the face of accountability (Hoffman and Patton 1997), and committees take actions to decrease the nonaudit service fee ratio in an effort to enhance both actual and perceived auditor independence (Abbott et al. 2003).

While it is prudent for audit committees to consider investors' perceptions, the fact that audit quality is sacrificed may not be in investors' best interest. This is particularly true if one considers that investors, when provided with "full" information, seem to prefer joint provision more than audit committees do. Based on an exploratory post-experimental question, we found that when asked to consider only required disclosures, i.e., to ignore information about the audit quality effects of joint provision, investors have preferences closer to those of public audit committees. This result seems consistent with Dopuch et al.'s (2003) findings that mandated fee disclosures may result in investors forming inaccurate beliefs of auditor independence. Together, our findings suggest that providing additional disclosure regarding a reason for the selection of the auditor for nonaudit services might be important given the auditors' independence is unobservable to investors (see, for example, Zhang 2003). Given one view that "mandatory disclosure [may not be] an effective means of redressing the inefficiency associated with the problem of [information] asymmetry" (Franco 2002, 240), it would seem important to examine the possible effects of voluntary disclosure.

By examining the effects of legislative mandates on experienced corporate directors' perceptions and decisions, this study addresses the appeals for empirical evidence on this issue. Results of this study may also assist regulators in evaluating the potential consequences of new policies and in assessing whether regulations achieve the intended goals of increasing audit and financial reporting quality. In addition, this paper contributes to existing literature on disclosure and financial reporting transparency. Prior research argues that managers sometimes make decisions to avoid increased financial reporting transparency (e.g., see, Healy and Palepu 2001; Hodge et al. 2004). This paper provides evidence that audit committees' decisions are affected by whether or not those decisions are publicly disclosed.

This paper also contributes to the growing body of research in the areas of corporate governance, audit committees, and auditor independence and suggests several avenues for future research. First, our investor-participants are clearly more sophisticated than the average investor. One extension would be to examine unsophisticated investors' preferences relative to both audit committee recommendations and sophisticated investor preferences. The findings in Hodge et al. (2004) suggest that unsophisticated investors—perhaps more so than sophisticated investors—would be more sensitive to fee disclosures in that "professional" audit committees are publicly linking fees to auditor independence. This may lead unsophisticated investors to place more emphasis on the fee-independence relation than they would otherwise.

Second, it would be interesting to examine the *types* of risks audit committees view as associated with fee disclosures. One typically associates public disclosures with increased legal liability; however, at a recent corporate governance conference, one legal scholar suggested that audit committees were considerably less concerned with disclosure-related legal liability and more troubled about reputation risk. Third, we posit that audit committees' recommendations are affected by required disclosures. This says, in effect, that committees differentially weight various factors relevant to their decisions. While we did not devise our research to specifically study decision models or the weights directors place on different pieces of information used to make their recommendations, future research along the lines of Peecher (1996) would be a fruitful avenue.

This study acknowledges all of the usual and expected limitations of experimental work. However, we believe by examining decisions of actual corporate directors, we limit some of the concerns about generalizability and perhaps provide a more complete picture of audit committee decisions when examined in conjunction with archival studies. While several archival studies explore the effects of joint provision by looking at proxies of audit quality, such as their relations with financial restatements (e.g., Kinney et al. 2004), abnormal accruals (e.g., Frankel et al. 2002), and going-concern opinions (e.g., DeFond et al. 2002), no prior studies have looked at the possible direct effects on audit committees' decisions.

REFERENCES

Abbott, L. J., S. Parker, G. F. Peters, and K. Raghunandan. 2003. An empirical investigation of audit fees, nonaudit fees, and audit committees. *Contemporary Accounting Research* 20 (Summer): 215–234.

Braiotta, L. 2004. *The Audit Committee Handbook.* New York, NY: John Wiley & Sons.

Bryan-Low, C. 2003. Accounting firms earn more from consulting—Despite SEC crackdown, buying nonaudit services by companies stays high. *Wall Street Journal* (April 19).

Copeland, J. E. 2002. Context of comments provided by Mr. Copeland to SEC roundtable on financial disclosure and auditor oversight. U.S. Securities and Exchange Commission, March 4. Washington, D.C.: SEC.

DeFond, M., K. Raghunandan, and K. R. Subramanyam. 2002. Do nonaudit service fees impair auditor independence? Evidence from going-concern audit opinions. *Journal of Accounting Research* 40: 1247–1274.

Dopuch, N., R. R. King, and R. Schwartz. 2003. Independence in appearance and in fact: An experimental investigation. *Contemporary Accounting Research* 20 (Spring): 79–114.

Federal Register. 2003. Strengthening the Commission's requirements regarding auditor independence—Part 2 of 2. Federal Register 68, No. 24. Washington, D.C.: Government Printing Office.

Ferguson, M. J., G. S. Seow, and D. Young. 2004. Nonaudit services and earnings management: UK evidence. *Contemporary Accounting Research* 21 (Winter): 813–842.

Flaming, L. 2005. Do nonaudit services affect investor judgments: Evidence using auditor fee proxy disclosures? Working paper, Monmouth University.

Franco, J. A. 2002. Why antifraud prohibitions are not enough: The significance of opportunism, candor and signaling in the economic case for mandatory securities disclosure. *Columbia Business Law Review*: 223–362.

Frankel, R. M. Johnson, and K. Nelson. 2002. The relation between auditors' fees for nonaudit services and earnings management. *The Accounting Review* 77 (Supplement): 71–105.

Gehrke, W. 2006. Even if you're private—What you should know about Sarbanes-Oxley. Available at: http://www.softrax.com/pdf/soxforprivatecompanies_revrec.pdf.

Goldwasser, D. L. 2002. The accounting profession's regulatory dilemma. *CPA Journal* 72 (5): 8–11.

Hartman, T. E. 2004. United States: The impact of Sarbanes-Oxley on private companies. September 14. Available at: http://www.mondaq.com/i_article.asp?articleid=28359.

Healy, P. M., and K. G. Palepu. 2001. Information asymmetry, corporate disclosure, and the capital markets: A review of the empirical disclosure literature. *Journal of Accounting & Economics* 31: 405–440.

Hirst, D. E., and P. E. Hopkins. 1998. Comprehensive income reporting and analysts' valuation judgments. *Journal of Accounting Research* 36: 47–75.

Hodge, F., J. Kennedy, and L. Maines. 2004. Does search-facilitating technology improve the transparency of financial reporting? *The Accounting Review* 79 (July): 687–703.

Hoffman, V. B., and J. M. Patton. 1997. Accountability, the dilution effect, and conservatism in auditors' fraud judgments. *Journal of Accounting Research* 35 (Autumn): 227–237.

Karmel, R. S. 2005. Realizing the dream of William O. Douglas—The Securities and Exchange Commission takes charge of corporate governance. *Delaware Journal of Corporate Law* 30: 79–144.

Kinney, W. R. Jr., Z-V. Palmrose, and S. Scholtz. 2004. Auditor independence, nonaudit services, and restatements: Was the U.S. Government right? *Journal of Accounting Research* 42 (June): 561–588.

Lerner, J. S., and P. E. Tetlock. 1999. Accounting for the effects of accountability. *Psychological Bulletin* 125 (2): 255–275.

Libby, R., R. Bloomfield, and M. W. Nelson. 2002. Experimental research in financial accounting. *Accounting, Organizations and Society* (December): 775–810.

Maines, L. A., and L. S. McDaniel. 2000. Effects of comprehensive income characteristics on nonprofessional investors' judgments: The role of financial-statement presentation format. *The Accounting Review* 75 (April): 179–207.

McDonough, W. J. 2004. Testimony before the Subcommittee on Capital Markets, Insurance and Government-Sponsored Enterprises, Committee on Financial Services and United States House of Representatives. June 24. Washington, D.C.: Government Printing Office.

Nam, S. 2005. The impact of nonaudit services on capital markets. Working paper, New York University.

Peecher, M. E., and D. N. Kleinmuntz. 1991. Discussion of experimental evidence on the effects of accountability on auditor judgments. *Auditing: A Journal of Practice & Theory* 10: 96–114.

———. 1996. The influence of auditors' justification processes on their decision: A cognitive model and experimental evidence. *Journal of Accounting Research* 34 (Spring): 125–140.

———, and I. Solomon. 2001. Theory and experimentation in studies of audit judgments and decisions: Avoiding common research traps. *International Journal of Auditing* 5: 193–201.

Pennington, J., and B. R. Schlenker. 1999. Accountability for consequential decisions: Justifying ethical judgments to audiences. *Personality and Social Psychology Bulletin* 25: 1067–1081.

Pitt, H. L. 2002. Testimony Concerning the Corporate and Auditing Accountability, Responsibility, and Transparence act. Before the Committee of Financial Services United States House of Representatives. March 20. Washington, D.C.: Government Printing Office.

Pletz, J. 2002. Several companies with Austin, Texas ties not planning accounting changes. *Knight Ridder Tribune Business News* (July 16).

Plitch, P. 2002. Pricewaterhouse executive speaks out on accounting issues. *Dow Jones Business News* (November 21).

PricewaterhouseCoopers. 2005. Barometer surveys. June 21. Available at: http: //www.barometersurveys.com/production/barsurv.nsf/vwAllNewsByDocID/834B5ECEF3C79C685257026006F2DC7.

Public Oversight Board (POB), Panel on Audit Effectiveness. 2000. *Report and Recommendations.* August 31. Stamford, CT: POB.

Raghunandan. K., and D. V. Rama. 2003. Audit committee composition and shareholder actions: Evidence from voting on auditor ratification. *Auditing: A Journal of Practice & Theory* 22 (September): 253–263.

Rice, B. 2000. Comments at public hearing on proposed auditor independence rules. Securities and Exchange Commission. September 13. Pace University. Available at: http://www.sec.gov/rules/extra/audmin2.htm.

Rittenhouse, L. 2001. Sun shines on nonaudit consulting fees. *AIMR Advocate* (September/October).

Securities and Exchange Commission (SEC). 2000. *Proposed Rule: Revision of the Commission's Auditor Independence Requirements.* Release No. 33-7870. June 30. Washington, D.C.: Government Printing Office.

Seidenfeld, M. 2001. The psychology of accountability and political review of agency rules. *Duke Law Journal* 51: 1059–1095.

Shedlarz, D. 2002. Context of comments provided by Mr. Shedlarz to SEC Roundtable on Financial Disclosure and Auditor Oversight. U.S. Securities and Exchange Commission. March 4, Washington, D.C.: SEC.

Siegel-Jacobs, K., and J. F. Yates. 1996. Effects of procedural and outcome accountability on judgment quality. *Organizational Behavior and Human Decision Processes* 65 (January): 1–17.

Simunic, D. 1984. Auditing, consulting, and independence. *Journal of Accounting Research* 22: 679–702.

Tetlock, P. E. 1983. Accountability and complexity of thought. *Journal of Personality and Social Psychology: Attitudes and Social Cognition* 45: 74–83.

———, L. Skitka, and R. Boettger. 1989. Social and cognitive strategies of coping with accountability: Conformity, complexity, and bolstering. *Journal of Personality and Social Psychology: Interpersonal Relations and Group Dynamics* 57: 632–641.

———, and R. Boettger. 1994. Accountability amplifies the status quo effect when change creates victims. *Journal of Behavioral Decision Making* 7: 1–23.

Turley, J. S. 2004. Testimony of James S. Turley, chairman and CEO Ernst & Young LLP, before the U.S. Senate Committee on Banking, Housing and Urban Affairs. September 9.

Turner, L. 2001. Audit committees: A roadmap for establishing accountability. Remarks to Conference Sponsored by Washington University School of Law and the Institute for Law and Economic Policy, Scottsdale, AZ, March 10.

———. 2002. Economic reality in financial reporting: A time for change. Speech to Australian Investment Management Conference, Sydney, Australia. September 23. Available at: http://www.cfawebcasts.org/cpe/what.cfm?test_id=304.

Unger, L. S. 2001. This year's proxy season: Sunlight shines on auditor independence and executive compensation. Speech of Acting SEC Chairman, U.S. Securities and Exchange Commission. June 25. Washington, D.C.: SEC.

U.S. House of Representatives. 2002. The Sarbanes-Oxley Act of 2002. Public Law 107-204 [H. R. 3763]. Washington, D.C.: Government Printing Office.

Weil, J., and J. Tannebaum. 2001. Big companies pay audit firms more for additional services. *Wall Street Journal* (April 10): C1.

Whisenant, S., S. Sankaraguruswamy, and K. Raghunandan. 2003. Evidence on the joint determination of audit and nonaudit fees. *Journal of Accounting Research* 41: 721–744.

Zhang, P. 2003. Discussion of: Independence in appearance and in fact: An experimental investigation. *Contemporary Accounting Research* 20 (Spring): 115–119.

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Performance Measure Properties and Delegation

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ABSTRACT: In this paper, I extend the organizational design literature by examining how the delegation choice is affected by the ability to resolve the incentive problem caused by this delegation. Based on the seminal papers by Grossman and Hart (1986) and Holmstrom and Milgrom (1994), I argue that the ability to resolve the incentive problem depends on the contractibility of financial performance measures versus nonfinancial performance measures, where the contractibility depends on the performance measure properties sensitivity, precision, and verifiability. The empirical results show that, if financial performance measures are "good" ("poor") incentive measures, i.e., high (low) on sensitivity, precision, and verifiability, then using these measures for incentive purposes increases (decreases) delegation. Overall, the results are consistent with the argument that firms design their decision-making process around the quality of contractible performance measures.

Keywords: *delegation; performance measure properties; contractibility; financial vs. nonfinancial performance measures; simultaneity.*

Data Availability: *Contact the author.*

I. INTRODUCTION

In this paper, I examine how the quality of contractible performance measures affects the delegation choice. There is a large stream of literature on the design of incentive systems and the design of these systems for lower-level managers has received increased attention in the empirical accounting literature. For example, Bushman et al. (1995) and Keating (1997) examine the factors associated with the importance of specific types of performance measures used to reward lower-level managers. Although these studies provide valuable insights into the incentives provided to lower-level managers, none of these studies incorporate the link between delegation and incentives. This omission is likely to be problematic given that the delegation choice is one of the crucial organizational design variables and affects the extent to which incentives need to be provided.

This paper is based on my dissertation at Maastricht University. I gratefully appreciate the comments made by my thesis advisors Willem Buijink, Frank Hartmann, and Ken Merchant. I am further grateful to Erik Peek, Madhav Rajan (editor), Mike Shields, Wim Van der Stede, Sally Widener, Jerry Zimmerman, two anonymous reviewers, seminar participants at Maastricht University, Nijenrode University, University of Amsterdam, University of Southern California, the ROA Workshop on Human Capital in Maastricht, and the 2005 AAA Management Accounting Section Conference for their comments and suggestions, and I am especially grateful to Chris Ittner.

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Empirical studies that address the determinants of delegation are rare. Baiman et al. (1995) examine the impact of the relative expertise of the principal and the relative importance of the business unit on the decision rights allocated to the business unit manager, but they ignore the effect of incentives. Recent studies by Nagar (2002), Abernethy et al. (2004), and Demers et al. (2004) do address the effect of performance measurement and incentives on delegation, but generally fail to find such an effect. For example, Nagar (2002) finds that incentive-based pay, which in his setting is predominantly based on earnings, does not affect delegation, although delegation does increase incentive-based pay. Similarly, Abernethy et al. (2004) find that the relative use of financial performance measures for incentive purposes does not affect the extent of delegation, but delegation does increase the emphasis on these measures. These results seem to suggest that once we delegate, we resolve the incentive problem by emphasizing financial measures for incentive purposes, but being able to resolve the incentive problem by emphasizing financial measures does not allow us to delegate. From a theoretical perspective, this conclusion is problematic, since the argument for the former relationship assumes the existence of the latter.

In this paper, I extend the organizational design literature by examining how the delegation choice is affected by the ability to resolve the incentive problem caused by this delegation. Based on the seminal papers by Grossman and Hart (1986) and Holmstrom and Milgrom (1994), I argue that the ability to resolve the incentive problem depends on the quality of contractible performance measures and, more specifically, on the quality of performance measures that aggregate information about all actions. An aggregate performance measure can complement the delegation choice because it provides the agent with the required discretion and the principal with an instrument to constrain the extraction of private benefits (Prendergast 2002). Comprehensive financial (accounting) measures are the most aggregate performance measures because they ultimately reflect the consequences of all decisions. However, the degree to which this aggregation feature can be exploited depends on the costs of contracting on financial measures. The more sensitive, precise, and verifiable financial measures are relative to other (nonfinancial) measures, the less costly it is to contract on financial measures. As a result, the better the relative incentive properties (quality) of financial performance measures, the greater their relative incentive use can complement the delegation choice, and thus the greater the extent of delegation.

I measure delegation as the day-to-day allocation of decision rights and the reason for using this particular measure is threefold. First, this measure of delegation allows for within-firm variation, which is necessary given the data limitations of using multiple respondents within a firm. Second, this measure of delegation has been used by Nagar (2002), Abernethy et al. (2004), and Demers et al. (2004) and allows me to tie my results back to these studies. Last, but not least, in contrast to more structural ways to implement delegation, such as divisionalization, the allocation of decision rights is the aspect of delegation that is most likely to be affected by incentive choices.

This paper contributes to the literature in the following ways. First, it empirically shows how the quality of contractible performance measures affects the delegation choice. More specifically, it shows that the relative incentive use of aggregate financial performance measures increases delegation when these measures have good incentive properties, which provides evidence that financial measures do play a role in the decision to delegate. Second, the results provide a plausible explanation for why previous research has been unable to find an effect of performance measurement and incentives on delegation, since these studies ignore the properties underlying the performance measures used for incentive purposes. Finally, it extends a small, but growing literature on the determinants of organizational design choices and enhances our knowledge of the interrelation between these choices.

The data that I use in the empirical analysis consist of both survey data and archival data, the latter of which is used to validate the use of performance measures for incentive purposes. I define the use of performance measures for incentive purposes broadly in this paper. That is, it reflects how important the performance measures are for periodic evaluations, salary increases, annual bonuses, and promotion possibilities. I choose these aspects because they match the type of incentives provided to the managers participating in this study. The performance measure properties that I examine relate to the extent to which a performance measure is influenced by (1) the manager's actions (sensitivity), (2) factors outside the control of the manager (precision), and (3) the measurement process (verifiability).

I find that, after controlling for the marginal benefits of delegation, the greater the relative sensitivity and precision of financial performance measures, the greater the returns to delegation when using these measures for incentive purposes. I further find that the greater the relative verifiability of financial performance measures, the greater the extent of delegation. Additional tests indicate that if financial performance measures are "poor" incentive measures, i.e., low on sensitivity, precision, and verifiability, then using these measures for incentive purposes creates an incentive-related cost of delegation and thus decreases delegation (cf., Nagar 2002). In contrast, if financial measures are "good" incentive measures, then using these measures for incentive purposes complements delegation choice and thus increases delegation (cf., Abernethy et al. 2004). These results are consistent with the argument that firms design their decision-making process around the quality of contractible performance measures (Prendergast 2002; Holmstrom and Milgrom 1994; Grossman and Hart 1986). When it is difficult or costly to contract on financial performance, firms reduce the need for incentive contracting by lowering the extent of delegation. Finally, robustness checks corroborate the above findings.

The remainder of this paper is organized as follows. In Section II, I present the theoretical analysis and hypothesis development. In Section III, I describe the sample selection and data collection and in Section IV the variable measurement and empirical specification. I discuss the empirical results in Section V. Finally, I provide a summary and conclusion in Section VI.

II. THEORY

Delegation and the Quality of Contractible Performance Measures

Organizational design, including the allocation of decision rights and the design of incentive systems, is the basic problem addressed by agency theory. In general, agency theory analyzes the situation in which a principal delegates authority to an agent and designs an incentive contract to motivate this risk and work-averse agent. The principal delegates authority to the agent because the agent has better decision-relevant information that is too costly to transfer due to, for example, the environmental uncertainty (Jensen and Meckling 1992). By doing this, the principal inevitably creates an incentive problem, which needs to be resolved through incentive contracting (Holmstrom 1979). Both Grossman and Hart (1986) and Holmstrom and Milgrom (1994) indicate that the design of firms reflect these contracting problems. These models demonstrate that when it is difficult or costly to contract on performance, firms arrange their affairs so as to reduce the need for incentive contracting (see also, Holmstrom 1999). The optimal allocation of decision rights is therefore affected by the ability to resolve the incentive problem caused by this allocation (Hubbard 2000).

To provide a simple formalization of the above discussion, I specify the value of delegation V as follows (cf., Baker and Hubbard 2003):

$$V = \underline{V} + ms - M(s,\sigma) \tag{1}$$

where $\underline{V}$ is a fixed quantity, s is the scope of the manager's activities due to the allocation of decision rights, m is the marginal benefit of allocating decision rights, σ is the ability to resolve the incentive problem, and $M(s,\sigma)$ is agency costs. I assume that $M_1 > 0$, $M_{11} > 0$, $M_2 < 0$, and $M_{12} < 0$, i.e., agency costs are increasing and convex in delegation, agency costs are decreasing in the ability to resolve the incentive problem, and the increased agency costs associated with delegation are lower the greater the ability to contract.

The optimal allocation of decision rights is then determined by setting the marginal costs of this allocation equal to the marginal revenues, i.e.:

$$m = M_1(s^*,\sigma). \tag{2}$$

Given the above assumptions, Equation (2) is invertible and the optimal allocation of decision rights is characterized by:

$$s^* = \phi(m,\sigma). \tag{3}$$

Equation (3) indicates that the optimal allocation of decision rights increases with increases in (a) the marginal benefits of delegation and (b) the ability to resolve the incentive problem caused by delegation. In this paper, I argue that the incentive properties (or quality) of some performance measures give these measures inherently greater potential to resolve the delegation-incentive problem, but the degree to which this potential can be exploited depends upon the cost to contract on these measures relative to other measures.

Delegation and Financial versus Nonfinancial Performance Measures

Broadening the scope of an agent's activities by delegating more decision rights provides the agent with substantial degrees of freedom to make trade-offs among these activities (Prendergast 2002; Jensen 2001). This creates a need for performance measures that allow for (more) discretion, but at the same time creates a need for constraining the agent's actions to prevent the extraction of private benefits.

Prendergast (2002) argues that this delegation-incentive problem can be addressed by tying pay to an aggregate measure of performance. "Aggregate" performance measures are measures that provide (some) information about "all" actions, while "specific" performance measures are measures that provide (some) information about a subset of actions. The use of an aggregate measure provides incentives for the agent to make trade-offs among all available activities, supporting the delegation of decision rights (Prendergast 2002). The aggregate measure also allows the principal to constrain the agent's actions to those in the principal's interest by tying this measure to pay. As a result, an aggregate performance measure can complement the delegation choice by allowing discretion to those with decision-relevant information (the agent), while also providing those lacking this information (the principal) with an instrument to constrain the extraction of private benefits (cf., Prendergast 2002).[1]

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[1] Two other reasons for why aggregate measures complement the delegation choice are the following. First, when broadening the scope of activities, it is simply too costly to have performance measures for each type of activity (Ittner and Larcker 2002; Banker and Datar 1989; Holmstrom and Milgrom 1987), which results in an increased preference for aggregate measures. Second, aggregation complements delegation as it allows the firm to economize on bounded rationality (Arya et al. 2004; Williamson 1975).

In contrast, using specific measures for subsets of the agent's available actions reduces the ability to make trade-offs, which effectively reduces delegation (Abernethy et al. 2004; Jensen 2001). Moreover, since the benefits of delegation are highest in settings where the principal has no idea what the agent should be doing, it is difficult to tie pay to multiple specific performance measures because the principal is unable to identify the specific actions the agent should take and, therefore, the specific measures and their weighting (Prendergast 2002).

Comprehensive financial (accounting) measures, such as net income or return on assets, represent the most aggregate performance measures because the full consequences of every action the agent takes ultimately flow through the financial statements. Various types of nonfinancial performance measures can also aggregate information about the agent's actions to some extent. For example, market-share provides information about all of the agent's customer acquisition decisions. Similarly, defect rates provide information about all of the agent's quality improvement activities. However, neither measure reflects the full consequences of these decisions, such as the costs of achieving market-share or defect goals or their revenue implications.[2] More importantly, they do not capture decisions other than customer acquisition and/or quality initiatives. In this respect, nonfinancial measures are "specific" performance measures, in the sense that they provide (some) information about a specific subset of actions (cf., Ittner and Larcker 1998; Fisher 1995; Wruck and Jensen 1994). This suggests that the principal will prefer the incentive use of comprehensive financial measures to the incentive use of more specific nonfinancial measures when delegation is greater.

However, the extent to which this aggregation feature of financial performance measures can be exploited for delegation purposes depends on the associated contracting costs. The contracting costs are driven by the incentive properties of the performance measures and in particular by the extent to which these measures are sensitive, precise, and verifiable. In general, the more precise and verifiable a measure, and the more sensitive it is to managerial actions, the greater the returns to using this measure for incentive purposes, *relative to* other measures (Feltham and Xie 1994; Banker and Datar 1989; Holmstrom 1979). The costs of contracting on aggregate financial measures can therefore be determined by using the specific nonfinancial measures as a benchmark, i.e., the more sensitive, precise, and verifiable financial measures are relative to nonfinancial measures, the less costly it is to contract on the financial measures.

As a result, the better the relative incentive properties (or quality) of financial performance measures, the greater the relative incentive use of these measures is able to address the incentive problem associated with delegation and, thus, the greater the extent of delegation.[3] Within this context, the ability to resolve the incentive problem σ in Equation (3) can therefore be written as:

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[2] As an example, an agent could dramatically drop the price and/or increase expenditures on advertising and customer service to increase market-share. Although this increase in market-share makes this measure informative about the two decisions, it does not reflect the full revenue and cost consequences of these decisions.

[3] I focus on the *relative* incentive use for the following reason. The argument is that aggregation of information for incentive purposes has the potential to complement delegation. Even though a higher absolute use of financial measures implies a higher absolute use of aggregate measures, this does not necessarily imply more aggregation for incentive purposes. In a given period, the absolute use of nonfinancial measures could be as high or even higher, which results in an emphasis on more specific measures and, thus, less overall aggregation. The relative incentive use of financial measures more accurately reflects the focus on overall aggregation and is also consistent with Abernethy et al. (2004) and Demers et al. (2004).

$$\sigma = \frac{INC_{FIN}}{INC_{NONFIN}} \cdot \frac{QUAL_{FIN}}{QUAL_{NONFIN}} \tag{4}$$

and the optimal delegation of decision rights s^* as:

$$s^* = \phi\left(m, \frac{INC_{FIN}}{INC_{NONFIN}} \cdot \frac{QUAL_{FIN}}{QUAL_{NONFIN}}\right) \tag{5}$$

where INC_{FIN} (INC_{NONFIN}) is the incentive use of financial (nonfinancial) performance measures and $QUAL_{FIN}$ ($QUAL_{NONFIN}$) is the quality of financial (nonfinancial) performance measures, as determined by the incentive properties. Equation (5) shows that the impact of the relative incentive use of financial performance measures on delegation is conditional on the relative quality of these measures.[4] Based on the above discussion, I state the following hypothesis.

H1: The impact of the relative incentive use of financial performance measures on the extent of delegation is increasing in the relative quality of these measures, as reflected by their relative incentive properties.

III. SAMPLE AND DATA

To test the prediction stated in the previous section, I gathered data from multiple sources. The primary data are survey questionnaire data, while the secondary data are interview data and proprietary archival data. Survey data are available for 105 managers employed in 6 firms. This section presents details on the sample selection and data collection.

Sample of Firms

This paper focuses on how the incentive use and properties of different types of performance measures affect delegation. In order to select the firms of interest to this study, I used the following procedure. I contacted Hay Management Consultants (HMC) and explained the research question of this study. HMC provided a list of clients that were likely to be eligible for this study, which contained the name of the firm, name of the Human Resource Manager, firm address, and telephone number.

I called the Human Resource (HR) managers of 50 firms to explain the research study and solicit their participation. Participation, at first, meant giving an interview. Out of the

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[4] More specifically, Equations (3) and (5) imply that the following is true:

$$\frac{\partial s^*}{\partial\left(\frac{INC_{FIN}}{INC_{NONFIN}}\right)} = \phi_2(\cdot)\left(\frac{QUAL_{FIN}}{QUAL_{NONFIN}}\right) = -\frac{M_{12}(\cdot)}{M_{11}(\cdot)}\left(\frac{QUAL_{FIN}}{QUAL_{NONFIN}}\right) > 0.$$

I assume that the agency cost function M(·) is such that the cross-partial:

$$\frac{\partial^2 s^*}{\partial\left(\frac{INC_{FIN}}{INC_{NONFIN}}\right)\partial\left(\frac{QUAL_{FIN}}{QUAL_{NONFIN}}\right)}$$

is positive.

50 firms, 15 (30 percent) agreed to an interview. Noteworthy is that the majority of the firms (17) that did not agree to participate were firms that were currently implementing a new or "updated" incentive system, which they characterized as "politically sensitive." Other reasons for nonparticipation were time constraints (11), reorganizations (5), and mergers (2).

The interview with the HR managers served two purposes. First, the interview was used to get a better understanding of the incentive system in place. It was important to examine if performance evaluation and compensation were a real issue and to what extent incentive compensation was important. This facilitated a more specific selection of firms relevant to this study. Second, the interview was used to ask the HR manager to support the research and to be its "champion" within the firm.

During the interviews, the content of participation was discussed in more depth. I explained that participation implied (1) the selection of respondents, (2) the distribution of questionnaires by internal mail, (3) the attachment of a letter of endorsement by the HR manager or a higher-level official, (4) administrative support for follow-up procedures, and (5) the provision of a detailed description of the respondents' annual bonus plan. Of the 15 firms interviewed, six declined further collaboration. The reasons why these firms did not participate were that they did not allow me to go "into the firm" (four firms) or that they did not have an incentive system (two firms). Of the nine firms that agreed to participate, three requested a more tailor-made study. To assure uniformity of the research design, I decided to design a separate study for each of these firms and to exclude them from the current study. As a result, the final sample consists of six firms. For completeness, it should be noted that none of these six firms have a long-term incentive plan, stock option plan, or any other equity ownership plan. In general, the firms provide incentives to their managers through periodic evaluations, salary increases, annual bonuses, and promotion possibilities. Descriptive statistics of the participating firms are provided in Table 1. The activities of the firms relate to the provision of services, trading, and production. Firm size, measured by the number of employees, ranges from 354 to 12,207.

Sample of Respondents

After the firms agreed to participate in the study, the HR manager of each firm selected the sample of respondents. In making the selection, the HR managers were given three criteria on which they should base their selection. First, the respondents should have managerial responsibilities, either as head of a functional department or as head of a division, business unit or something similar. To assure a minimum level of managerial responsibilities, the respondent's job design should have a weight greater than 400 Hay-points.[5] Second, the respondent should have an annual bonus plan. Finally, the respondent should have experienced at least one annual performance evaluation cycle.

The HR managers were specifically asked to select as many respondents as possible and as diverse as possible within the above three constraints. Within-firm variance in the two main choice variables in this study, i.e., extent of delegation and the relative incentive use of financial performance measures, should be observed for the following reason. The same performance measure in the same firm is unlikely to have the same incentive properties for different groups of employees within the firm. A diverse selection of respondents should therefore create a sample of employees within a firm who contribute differently to

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[5] The Hay-points are based on the Hay Guide Chart Profile Method, a system that compares the value of jobs based on multiple factors such as accountability and know-how (Flannery et al. 1996, 20). A weight greater than 400 relates to higher-level personnel.

TABLE 1
Descriptive Statistics of Participating Firms
(n = 6)

Firm	Activity	# of Employees	Sample
A	Provision of financial services	1,690	9
B	Trade of machinery and provision of technical services	12,207	28
C	Provision of life and indemnity insurance	1,275	14
D	Trade of pharmaceutical products	401	31
E	Production of food products	7,482	5
F	Financial leasing	354	18

(the same type of measure of) performance and are also exposed to different externalities. Furthermore, within a hierarchy, it is each supervisor's individual decision how much emphasis to put on what measure for incentive purposes (broadly defined) and how much of the day-to-day decisions to delegate, though obviously within certain boundaries. These differences and (relative) freedom of choice should create within-firm variance in the variables of interest. The selection by the HR managers resulted in 202 managers who were asked to participate in the study.

Questionnaire Design

To maximize the response rate, I designed the questionnaire according to the guidelines of Dillman's (1978) *Total Design Method.* I pre-tested the questionnaire with seven academics and all six HR managers, which resulted in minor revisions of the questionnaire. Furthermore, I administered a follow-up by telephone and mail. Of the 202 questionnaires that were distributed, 114 were returned. This corresponds to an overall response rate of 56 percent. Of the 114 questionnaires returned, nine have missing data and the final sample therefore consists of 105 observations.

As the response rate is not 100 percent, though satisfactory, I conduct a test for nonresponse bias. I split the sample at the firm-specific median response time and calculate a t-test for differences in means for each of the variables of interest to this study. The results indicate that there are no significant differences between early and late respondents for any of the variables. This suggests the absence of nonresponse bias.

For the two main choice variables in this study, i.e., extent of delegation and the relative incentive use of financial performance measures, the within-firm variance as a percentage of total variance equals 72 percent and 73 percent, respectively. Furthermore, the within-firm variance for delegation (relative incentive use of financial performance measures) ranges from 70–77 percent (70–79 percent) after piece-wise deletion of each firm, which indicates that this variance is not driven by a single firm.[6] As a result, the selection of respondents was successful in terms of creating within-firm variance in the sample.

Proprietary Archival Data

Of the six firms, five firms provided archival data with respect to the annual bonus contract of the respondents. This data is used to validate the survey-based measures of performance measure use. Four firms have bonus contracts for their managers that are uniform in terms of the percentage of the bonus that is dependent on financial (nonfinancial) performance, although the number of performance measures in each category can vary significantly across managers.[7] However, I have no data with respect to the number of performance measures for each of the respondents in these four firms. Two firms have tailor-made manager-specific contracts, but only one firm provided that data.[8] As a result, archival data are available for 87 of the 105 managers for five out of six firms. The sample used to validate the survey-based measure of incentive use of performance measures is therefore a (large) subset of the full sample used to test the hypothesis.

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[6] Similar results apply for the piece-wise deletion of groups of two firms.

IV. VARIABLE MEASUREMENT AND EMPIRICAL SPECIFICATION

In this section, I describe the measurement of the variables used in the empirical analysis and the empirical specification used to test the prediction stated in Section II. All measurement instruments are presented in Appendix A.

Delegation

I measure delegation of authority by taking the scale used by Gordon and Narayanan (1984), which is similar to that used by Nagar (2002) and Abernethy et al. (2004). The instrument asks the respondents to indicate the extent to which they have decision-making authority with respect to (1) development of new products and services, (2) hiring and firing of personnel, (3) selection of large investments, (4) budget allocations, and (5) pricing decisions. A five-point, fully anchored scale is used to indicate the level of authority. Principal component analysis reveals one factor with an eigenvalue greater than 1, which explains 51 percent of the total variance. The factor loadings of the five items range from 0.58 to 0.82. Given that all five items appear to measure one underlying construct, I compute the *DELEGATION* construct by summing and averaging the standardized scores of the five items (Cronbach's $\alpha = 0.76$).[9]

Relative Incentive Use of Financial Performance Measures

I measure the relative use of financial performance measures for incentive purposes by comparing the importance of financial measures for evaluating and compensating the managers to that of nonfinancial measures. In order to measure this construct and the related performance measure properties as accurate as possible, I focus on three types of performance measures: (1) financial, (2) internal nonfinancial, and (3) external nonfinancial. This classification is partly based on prior literature (e.g., Ittner et al. 1997; Larcker 1981) and partly on the interviews with the human resource managers of the participating firms. First, the financial performance measures are defined as the "traditional" aggregate financial performance measures, such as return on assets and net income.[10] Second, the internal nonfinancial performance measures consist of nonfinancial measures that are directly related to the tasks performed. Examples of these performance measures are productivity, efficiency, and successful implementation of projects. Finally, the external nonfinancial performance measures are defined as those nonfinancial measures that reflect performance in the market. Examples of these performance measures are customer satisfaction, market share, and market growth.

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[7] As an example, one of the four firms provided me, on request, the opportunity to look into a number of manager-specific contracts. For all these managers, 43.75 percent of the bonus was based on nonfinancial performance measures, but the number of nonfinancial performance measures for each manager ranged from 4 to 12. Given the incentives for a particular performance category, increasing the number of performance measures within that category can lead to a dilution of incentives if managers need to spread their effort over too many measures. As a result, the number of performance measures significantly affects the actual level of incentives provided to managers.

[8] The archival data can be linked to the questionnaire data through codes. Each questionnaire contains a unique code that is printed on the final page of the questionnaire. The HR manager provides the archival data mentioning the codes instead of the respondent's name, which does not violate the anonymity principle.

[9] In computing all constructs, I use unit-weighted average standardized scores because these have preferred psychometric properties relative to regression estimates of factor scores (Grice and Harris 1998).

[10] All examples presented with respect to each type of performance measure are measures actually used for incentive purposes by the participating firms.

For each type of performance measure, I use eight items to measure its importance for (1) evaluation purposes, (2) monetary compensation, and (3) nonmonetary rewards. I use these specific items because they are in accordance with the actual type of incentives provided by the participating firms. A five-point, fully anchored scale is used to indicate the importance of each type of performance measure for incentive purposes. To measure the relative incentive use, I calculate for each item the difference between the scores for financial performance measures and the two nonfinancial performance measures by subtracting from the financial performance measure response both of the responses for the nonfinancial measures.

Principal component analysis of the relative measures of the eight items reveals one factor with an eigenvalue greater than 1, explaining 65 percent of the total variance. The factor loadings of the eight items range from 0.73 to 0.88. Based on these results, I compute the construct *REL_INC_FPM* by summing and averaging the standardized scores of the eight items (Cronbach's $\alpha = 0.92$).

To validate this survey-based construct of the relative use of financial performance measure for incentive purposes, I test whether *REL_INC_FPM* is consistent with the relative weight on financial performance measures (%FPM) stated in the annual bonus contract. Despite the fact that the annual bonus is but one of the many types of incentives provided to the managers in this study, the information contained in the annual bonus systems should be sufficient to allow for a validation of the survey-based measure. I split the sample into two groups of approximately equal size based on the median reported %FPM. I specifically test whether the mean within each subsample is significantly different from zero and whether the means between subsamples are significantly different. The results indicate that in the "low-%FPM" subsample, the mean relative incentive use of financial performance measures is significantly negative ($p < 0.05$; two-tailed), which implies that the relative incentive use of financial performance measures in this subsample is, on average, significantly lower than in the full sample. The opposite applies to the "high-%FPM" subsample, in which the mean relative incentive use of financial performance measures is significantly positive ($p < 0.01$; two-tailed). Further, the means of the two subsamples are significantly different from each other ($p < 0.01$; two-tailed). These results provide some evidence of the validity of the survey-based measure.

Performance Measure Properties

To measure the performance measure properties, I develop new constructs. The survey questionnaire contains, for both financial and the two nonfinancial performance measures, 15 statements concerning the performance measure properties. A five-point, fully anchored scale is used to indicate the level of agreement with these statements. The 15 statements, listed in Table 2, relate to the extent to which each type of performance measure is influenced by (1) the manager's actions, (2) factors outside the control of the manager, and (3) the measurement process. For each item, I calculate the difference between the scores for financial performance measures and the two nonfinancial performance measures by subtracting from the financial performance measure response both of the responses for the nonfinancial measures.

TABLE 2

Principal Component Analysis with Oblique Rotation of the 15 Items Related to the Properties of Financial Performance Measures Relative to those of Nonfinancial Performance Measures

(only factors loadings > 0.45 are shown; n = 105)

Items	Relative Sensitivity FPM	Relative Precision FPM	Relative Verifiability FPM
My performance expressed in financial performance measures is strongly affected by			
a. ... changes in economic conditions		0.74	
b. ... decisions made in other parts of the organization			
c. ... changes in the behavior of customers		0.75	
d. ... changes in the behavior or strategies of suppliers		0.73	
e. ... changes in the behavior or strategies of competitors		0.66	
f. Whether I function well or not as a manager can be expressed accurately in financial performance	0.48		
g. Many of the activities and tasks that I perform do not show up in financial performance[a]	0.64		
h. If I perform well as a manager, it is directly reflected in better financial performance	0.79		
i. Working hard leads to better financial performance	0.77		
j. Devotion and effort in my job leads to better financial performance	0.78		
k. Providing effort in my job leads to better financial performance	0.79		
l. The measurement of financial performance is objective and verifiable			0.76
m. The measurement of financial performance is done by objective persons			0.91
n. An independent person verifies the measurement of financial performance			0.88
o. The measurement of financial performance is predominantly of a quantitative nature			
Eigenvalue	3.66	2.45	1.91

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[a] This item is reverse-coded.

Principal component analysis of the relative measures of the 15 items reveals three independent factors with eigenvalues greater than 1, explaining 53 percent of the total variance. The factor loadings after applying oblique rotation are shown in Table 2. Examining the factor loadings reveals that the items loading on factor 1 relate to the impact the manager has on performance, the items loading on factor 2 relate to the impact of uncontrollable factors, while the items loading on factor 3 relate to the objectivity and verifiability of the measure. Given the results of the principal component analysis, I use the following three performance measure properties in the empirical analysis: (1) relative sensitivity of financial performance measures (*REL_SEN_FPM*), (2) relative precision of financial performance measures (*REL_PREC_FPM*), and (3) relative verifiability of financial performance measures (*REL_VERIF_FPM*). I measure the constructs *REL_SEN_FPM*, *REL_PREC_FPM*, and *REL_VERIF_FPM* by summing and averaging the standardized scores of the items that load on respectively factor 1, factor 2, and factor 3. Cronbach's alphas for these measures are 0.81, 0.70, and 0.83, respectively.

Control Variables

As stated in Equation (3), the extent of delegation is also affected by the marginal benefits of delegation, which relate to the cost of transferring knowledge from lower levels in the organization to the top (Christie et al. 2003; Jensen and Meckling 1992). The greater the knowledge transfer cost the greater the delegation of authority. Both the external environment and the internal environment of an organization affect the knowledge transfer costs. The knowledge transfer costs are especially high in uncertain environments and in larger firms (Christie et al. 2003; Jensen and Meckling 1992). If environmental uncertainty increases, then there is a need to respond quickly to changes in the environment and, since it is too costly for top management to acquire the necessary information for this purpose, delegation increases (e.g., Nagar 2002). Furthermore, if the size of the unit is large, then the amount of information that needs to be processed and transferred is too large and therefore too costly, which results in increased delegation (e.g., Christie et al. 2003; Jensen and Meckling 1992). Finally, if there are multiple hierarchical levels, then it is costly to centralize decision making, since this would require direct communication with and the transfer of knowledge from each individual layer (Melumad et al. 1992). Although such a hierarchical structure can distribute the burden of information processing more evenly within the organization, the benefits of this are only received when decision rights are actually delegated (e.g., Melumad et al. 1992, 1995), which implies that delegation is higher when multiple hierarchical layers are present.[11] Based on prior theory, I include environmental uncertainty, size, and the number of hierarchical levels as control variables for the marginal benefits of delegation.

The environmental uncertainty variable is derived from the scales used by Govindarajan (1984) and Merchant (1990) and consists of five attributes with respect to the respondent's work environment. The five attributes relate to the behavior of (1) customers, (2) competitors, and (3) suppliers, as well as (4) technological developments and (5) political and/or legal changes. Ten items are used to indicate to what extent each of these five attributes is *predictable* and has on *impact* on the respondent's job and unit. A six-point, fully anchored scale is used, which consists of a five-point scale to indicate the extent of predictability and impact and an additional option that can be used to indicate that the specific factor is not part of the respondent's work environment.

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[11] The hierarchical structure is potentially an endogenous variable. However, such a structural decision is most likely to precede the choices of the incentive use of performance measures and the day-to-day allocation of decision rights, and I therefore assume that it is exogenous in the empirical analysis. The results presented in the next section are not sensitive to the inclusion of the proxy for the number of hierarchical levels.

Similar to the way in which Khandwalla (1972) and Libby and Waterhouse (1996) measure competition, I compute the environmental uncertainty variable by multiplying, for each attribute, the ratings on impact by predictability and taking the square root of the product.[12] Principal component analysis reveals two factors with an eigenvalue greater than 1 that explain 60 percent of the total variance. After oblique rotation of the factor solution, the following two indicators of environmental uncertainty are identified: (1) uncertainty related to the input-side of the firm (*ENV_UNC_IN*; attributes: technological developments and the behavior of suppliers) and (2) uncertainty related to the output-side of the firm (*ENV_UNC_OUT*; attributes: behavior of customers and competitors). As a result, I measure *ENV_UNC_IN* and *ENV_UNC_OUT* by summing and averaging the standardized scores of their respective two attributes (Cronbach's $\alpha = 0.62$ for both constructs).

I measure *SIZE* by the ratio of the number of employees within the manager's unit to the total number of employees. The number of hierarchical levels is proxied by the dummy variable *HIER_LEVELS*, which equals 1 if there are multiple hierarchical levels below the manager's unit, and 0 otherwise.

Descriptive statistics for all variables are provided in Table 3, while Table 4 presents the Pearson and Spearman correlations between the independent variables, none of which cause multicollinearity concerns. Table 4 shows that, among the performance measure properties, only *REL_SEN_FPM* and *REL_VERIF_FPM* are significantly correlated (positive). In addition, *REL_SEN_FPM* and *REL_VERIF_FPM* are significantly positively correlated

TABLE 3
Descriptive Statistics of Variables
(n = 105)

Variable	Mean	Std. Dev.	Actual Range	Cronbach's α
DELEGATION	0	0.71	−1.82–1.49	0.76
REL_INC_FPM	0	0.81	−2.50–2.15	0.92
REL_SEN_FPM	0	0.71	−1.63–1.94	0.81
REL_PREC_FPM	0	0.73	−2.31–1.65	0.70
REL_VERIF_FPM	0	0.86	−1.61–2.07	0.83
ENV_UNC_IN	0	0.86	−2.71–1.78	0.62
ENV_UNC_OUT	0	0.85	−4.04–1.91	0.62
SIZE	0.04	0.07	0.00–0.63	—
HIER_LEVELS (dummy)	0.32	0.47	—	—

See Appendix A for variable measurement and definitions.

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[12] The five-point scale for "impact" ranges from low impact (score of 1) to high impact (score of 5) and for "predictability" from high predictability (score of 1) to low predictability (score of 5). If the attribute is not relevant, then a score of 0 is attached to that attribute regarding its impact and predictability. For each of the five attributes, I multiply the score for impact by that for predictability, where higher scores for the product term reflect greater uncertainty and the scores range from 0 to 25. The logic underlying the multiplication is that given the impact of the attributes, increased unpredictability makes it more difficult to control for this impact, which increases uncertainty. Finally, for each attribute, I take the square root of the product term to revert back to a six-point scale (0–5).

TABLE 4
Correlation Coefficients between the Independent Variables
(n = 105)

Variable	1	2	3	4	5	6	7	8
1. *REL_INC_FPM*	1.00	0.53***	−0.00	0.31***	0.03	0.08	−0.10	0.35***
2. *REL_SEN_FPM*	0.48***	1.00	−0.05	0.30***	−0.05	−0.01	0.01	0.08
3. *REL_PREC_FPM*	0.03	−0.04	1.00	−0.08	0.21**	0.01	−0.11	0.07
4. *REL_VERIF_FPM*	0.32***	0.30***	−0.09	1.00	−0.03	0.10	−0.10	0.20**
5. *ENV_UNC_IN*	0.11	−0.03	0.17*	−0.03	1.00	0.17*	−0.16*	0.15
6. *ENV_UNC_OUT*	0.10	−0.06	−0.00	0.09	0.26***	1.00	0.04	0.21**
7. *SIZE*	0.08	0.11	−0.05	−0.07	−0.07	−0.09	1.00	0.02
8. *HIER_LEVELS*	0.36***	0.08	0.11	0.19**	0.16	0.22**	0.10	1.00

***, **, * Significant at the 1 percent, 5 percent, and 10 percent levels, respectively (two-tailed).
Pearson (Spearman) correlation coefficients are presented below (above) the diagonal.
See Appendix A for variable measurement and definitions.

with *REL_INC_FPM*. Finally, although the principal component analysis revealed two separate constructs for environmental uncertainty, these constructs are significantly positively correlated.

Empirical Specifications

Based on Equation (5), the hypothesis, and the above identification of variables, I estimate the following two equations to examine the determinants of the extent of delegation:

$$\begin{aligned} DELEGATION_{ij} = {} & \alpha_0 + \sum_{j=1}^{5}\alpha_j D_j + \alpha_6 ENV\_UNC\_IN_{ij} + \alpha_7 ENV\_UNC\_OUT_{ij} \\ & + \alpha_8 SIZE_{ij} + \alpha_9 HIER\_LEVELS_{ij} + \alpha_{10} REL\_SEN\_FPM_{ij} \\ & + \alpha_{11} REL\_PREC\_FPM_{ij} + \alpha_{12} REL\_VERIF\_FPM_{ij} \\ & + \alpha_{13} REL\_INC\_FPM_{ij} + \varepsilon^1_{ij} \end{aligned} \tag{6}$$

$$\begin{aligned} DELEGATION_{ij} = {} & \alpha_0 + \sum_{j=1}^{5}\alpha_j D_j + \alpha_6 ENV\_UNC\_IN_{ij} + \alpha_7 ENV\_UNC\_OUT_{ij} \\ & + \alpha_8 SIZE_{ij} + \alpha_9 HIER\_LEVELS_{ij} + \alpha_{10} REL\_SEN\_FPM_{ij} \\ & + \alpha_{11} REL\_PREC\_FPM_{ij} + \alpha_{12} REL\_VERIF\_FPM_{ij} \\ & + \alpha_{13} REL\_INC\_FPM_{ij} \\ & + \alpha_{14} REL\_INC\_FPM_{ij} \times REL\_SEN\_FPM_{ij} \\ & + \alpha_{15} REL\_INC\_FPM_{ij} \times REL\_PREC\_FPM_{ij} \\ & + \alpha_{16} REL\_INC\_FPM_{ij} \times REL\_VERIF\_FPM_{ij} + \varepsilon^2_{ij}. \end{aligned} \tag{7}$$

The variables relate to manager i in firm j and D_j is a dummy variable that is 1 if firm j, and 0 otherwise, in order to control for firm-level effects. Both Equations (6) and (7) estimate whether delegation is a function of environmental uncertainty (input- and output-related), size, number of hierarchical levels, and the relative sensitivity, precision, verifiability, and incentive use of financial performance measures. Equation (6) examines whether there exists a direct association between delegation and the relative incentive use of financial performance measures. This equation is, more or less, consistent with the empirical specifications used in previous studies (e.g., Abernethy et al. 2004; Demers et al. 2004) and is therefore estimated as a first step to tie my results back to these studies. Equation (7), on the other hand, explicitly tests the hypothesis that the association between delegation and the relative incentive use of financial performance measures is a positive function of the performance measure properties. That is, it examines whether financial performance measures with better incentive properties lead to greater returns to delegation when using these measures for incentive purposes. Hypothesis 1 specifically predicts that the coefficients for the interaction terms α_{14}, α_{15}, and α_{16} are positive.

Both Equations (6) and (7) are initially estimated using OLS. However, to control for the possibility that the choice variable "relative incentive use of financial performance measures" causes endogeneity within the structural model of interest, I also estimate both equations using two-stage least squares (TSLS). Given that Equation (7) examines interactions between the endogenous variable and multiple exogenous variables, I estimate this equation using the Heckman and Vytlacil (1998) two-stage least squares estimator (hereafter, HV-estimator). This method basically (1) estimates the endogenous variable (first-stage), (2) uses the predicted values to create the interaction terms, and then (3) uses the predicted values and created interactions in estimating Equation (7) (second-stage).[13]

To apply TSLS, it is necessary to have instruments, i.e., variables that are correlated with the explanatory variable (relevant) and uncorrelated with the structural error term (exogenous). I use two instrumental variables for the relative incentive use of financial performance measures that are intuitively appealing and for which specification tests indicate they are relevant and exogenous. The instruments that I use for the relative incentive use of financial performance measures focus on the extent to which managers are able to effectively communicate in financial (nonfinancial) terms because of their functional background. I expect that the relative incentive use of financial performance measures is higher (lower) for those managers who are (not) able to effectively communicate in financial terms. Specifically, I expect that managers with a functional background in sales are able to effectively communicate in financial terms, while managers with a functional background in R&D are much more able to effectively communicate in nonfinancial terms. Although functional background can affect the *type* of task that is *performed*, there is no reason to assume that it affects the *extent* of *delegation*. Thus, I expect functional background to satisfy both the relevance and exogeneity criterion.

To determine relevance, I test for the first-stage exclusion restriction and the partial R^2. The first-stage exclusion restriction tests the (joint) significance of adding the instruments to the reduced form equation, while the partial R^2 tests how much of the unexplained variance in the reduced form equation can be explained by adding the instruments (e.g., Wooldridge 2002; Bound et al. 1995). The results show that the test of the first-stage exclusion restriction is significant ($p < 0.01$) and the partial R^2 equals 10 percent, which provides evidence of relevance. To determine exogeneity of the instruments, I test for the over-identifying restrictions. This test regresses the residuals of the second-stage on *all* exogenous variables, where the R^2 of the model should be close to zero if the instruments are exogenous (e.g., Larcker and Rusticus 2004; Wooldridge 2002).[14] The test of the over-identifying restrictions indicates that the null-hypothesis of exogenous instruments cannot be rejected. The problem with this latter test is that it can have low power to detect endogeneity in small samples. I therefore perform a sensitivity analysis, which Larcker and Rusticus (2004) label as the "unconstrained second-stage."[15] The results of the sensitivity analysis (for Equation (6)) show that the coefficients for the instruments are very close to the TSLS estimate and not significantly different from each other ($p = 0.99$). Overall, the tests cannot reject the exogeneity of the instruments used. Based on these tests and the above-mentioned arguments, I use two dummy variables as instruments, i.e., *SALES_MAN* and *R&D_MAN*, which equal 1 if the manager's functional specialization is Sales (R&D), and 0 otherwise.

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[13] An alternative for the HV-estimator to estimate equations with interactions between endogenous and exogenous variables is the instrumental variable (IV) estimator (Wooldridge 2003). As stated, the HV-method uses the predicted values of the endogenous variable and the interactions created with these predicted values as *variables* in Equation (7). In contrast, the IV-method proposed by Wooldridge (2003) uses the predicted values of the endogenous variable and the interactions created with these predicted values as *instruments* and subsequently estimates Equation (7) using IV procedures. The difference between the HV-method and the IV-method is that the latter method does not require the linearity assumption in the first-stage reduced form equation to hold and differences can thus arise if this requirement does not hold. Although the regression specification error test (RESET) proposed by Ramsey (1969) cannot reject the null-hypothesis that the reduced form equation is correctly specified, I re-estimate Equation (7) using the method proposed by Wooldridge (2003). The results, not tabulated, are qualitatively similar to those reported in Table 5 and show that all coefficients have the same sign and approximately the same magnitude as the HV-estimators.

[14] Formally, nR^2 is distributed χ^2 with K-L degrees of freedom, where K is the number of instruments, L is the number of endogenous variables, and $K>L$.

[15] The unconstrained second-stage regresses the dependent variable on all exogenous variables, but where each exogenous variable is replaced by the product of its original value and its first-stage coefficient. The resulting coefficients for the instruments should be close to each other and, therefore, close to the TSLS estimate if the instruments are valid (Larcker and Rusticus 2004).

To test for statistical significance, I do not rely on normal theory standard errors, but rather use bias-corrected bootstrapped confidence intervals (Jeong and Maddala 1993). I use 2,000 iterations with random resampling to estimate the parameters and to compute the bias-corrected confidence for each parameter, which is used to determine the level of statistical significance.[16] The bootstrap method has the advantage that it requires fewer assumptions than traditional methods and is generally more accurate. It mitigates problems associated with normal theory standard errors due to, for example, sampling variation, nonnormality, and heteroscedasticity, which is particularly relevant for TSLS given its ambiguous finite sample properties.[17]

V. RESULTS

Performance Measure Properties and Delegation

Table 5 presents the results of the OLS and TSLS estimation of the determinants of delegation.[18] The results for Model 1_{OLS} show that *DELEGATION* is negatively affected by *REL_PREC_FPM*, positively affected by *REL_VERIF_FPM*, and not affected by *REL_SEN_FPM* and *REL_INC_FPM*. This latter finding suggests that the relative incentive use of financial performance measures does not affect delegation in a simple linear fashion, which is consistent with the findings of Abernethy et al. (2004) and Demers et al. (2004). It further suggests that the more precise financial measures are relative to nonfinancial measures, the lower the extent of delegation, the more verifiable the higher the extent of delegation, while sensitivity has no direct effect. Overall, these results are inconsistent with delegation being directly associated with the relative incentive use of financial performance measures.

Regarding the control variables for the marginal benefits of delegation, the results for Model 1_{OLS} show that *DELEGATION* is positively affected by *ENV_UNC_IN*, *SIZE*, and *HIER_LEVELS*. This suggests that more decision rights are delegated, the greater the environmental uncertainty on the input-side of the firm, the greater the size of the unit, and the more hierarchical layers below the manager. Finally, Model 1_{TSLS} provides similar results and I therefore refrain from discussing these in detail.

Columns four and five of Table 5 present the results of adding interaction terms between the relative incentive use of financial performance measures and the relative properties of these measures, which provides the empirical test of the hypothesis stated in Section II. The results for Model 2_{OLS} are similar to those for Model 2_{TSLS} with the notable exception of the interaction term between *REL_INC_FPM* and *REL_SEN_FPM*. Although the coefficient for the interaction term is positive in both models, the Durbin-Wu-Hausman test

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[16] The parameters of interest here are the OLS parameters in the OLS models and the second-stage parameters in the TSLS models.

[17] An additional benefit of the bootstrapping procedure is that hypothesis testing is less troubled by potential multicollinearity issues. Although it is not at all clear whether multicollinearity is an issue in regressions with interaction terms (see, e.g., Hartmann and Moers 1999, 302), the general consequences of severe multicollinearity are large standard errors and unstable regressors. The bootstrapped confidence interval does not rely on normal theory standard errors and does not rely on a single estimate, which makes hypothesis testing less sensitive to multicollinearity.

[18] Appendix B presents the results of the first-stage of the TSLS regression, i.e., the estimates of the reduced form equation for the relative incentive use of financial performance measures.

TABLE 5
OLS and Two-Stage Least Squares Estimation of the Impact of the Relative Incentive Use and Properties of Financial Performance Measures on the Extent of Delegation

(n = 105)

	DELEGATION			
Independent Variables	**Model 1_{OLS}**	**Model 1_{TSLS}**	**Model 2_{OLS}**	**Model 2_{TSLS}**
Intercept	−0.63***	−0.62***	−0.64***	−0.62***
Firm$_A$	0.95***	0.93***	0.92***	0.92***
Firm$_B$	0.60***	0.58*	0.56***	0.50†
Firm$_C$	0.66***	0.65**	0.67***	0.61**
Firm$_D$	0.22	0.22	0.19	0.18
Firm$_E$	0.44	0.40	0.47	0.36
ENV_UNC_IN	0.09*	0.10†	0.10*	0.09
ENV_UNC_OUT	−0.01	−0.01	−0.01	−0.02
SIZE	1.99**	1.95**	1.84**	1.77**
HIER_LEVELS	0.40**	0.40**	0.39**	0.38**
REL_SEN_FPM	−0.08	−0.10	−0.07	−0.10
REL_PREC_FPM	−0.11*	−0.11*	−0.07	−0.06
REL_VERIF_FPM	0.16**	0.16**	0.16**	0.15**
REL_INC_FPM	−0.07	−0.03	−0.03	0.05
REL_INC_FPM × *REL_SEN_FPM*#			0.11	0.20**
REL_INC_FPM × *REL_PREC_FPM*			0.15*	0.19*
REL_INC_FPM × *REL_VERIF_FPM*			0.06	0.00
Adjusted R^2	0.35***	0.34***	0.35***	0.35***
Incremental R^2	0.15***	0.15***	0.17***	0.17***
Partial R^2	0.21***	0.21***	0.24***	0.24***

***, **, *, †Significant at the 1 percent, 5 percent, 10 percent, and 15 percents levels, respectively (two-tailed test using bias-corrected bootstrapped confidence intervals).

#Indicates significant endogeneity at the 10 percent level based on the Durbin-Wu-Hausman test.

The incremental R^2 indicates the total variance explained by the independent variables incremental to firm effects. The Partial R^2 indicates how much of the within-firm variance is explained by the independent variables.

Model 2_{TSLS} is estimated using the Heckman and Vytlacil (1998) two-stage least squares estimator.

The instrumental variables for the endogenous variable *REL_INC_FPM* in Model 1_{TSLS} and Model 2_{TSLS} are *SALES_MAN* and *R&D_MAN*. The first-stage adjusted R^2 equals 44 percent and the partial R^2 equals 10 percent. The test of the first-stage exclusion restriction is significant ($p < 0.01$), while the test of over-identifying restrictions indicates that the null-hypothesis of exogenous instruments cannot be rejected ($p > 0.90$ in both models). This suggests that the instruments are relevant and exogenous.

See Appendix A for variable measurement and definitions.

indicates that the OLS coefficient has a significant downward bias.[19] I therefore prefer to interpret the TSLS results.

The results for Model 2_{TSLS} show that there is a positive direct effect of *REL_VERIF_FPM* and a positive interactive effect between *REL_INC_FPM* and respectively *REL_PREC_FPM* and *REL_SEN_FPM* on *DELEGATION*. The first finding is consistent with Model 1. In contrast, the latter findings suggest that the relative sensitivity and precision and the relative incentive use of financial performance measures positively reinforce each other in determining the extent of delegation. That is, the greater the relative sensitivity and precision of financial performance measures, the greater the returns to delegation when using these measures for incentive purposes.[20]

The main effect for *REL_INC_FPM* in the interactive regression is not statistically significant. This main effect represents the effect of *REL_INC_FPM* on *DELEGATION* when the values of all variables with which it interacts—i.e., the performance measure properties—equal zero (Hartmann and Moers 1999, 300). Given that the zero value of the performance measure properties represents the sample average (see footnote 9), this implies that the relative incentive use of financial performance measures does not affect delegation "on average," which is consistent with Abernethy et al. (2004) and Demers et al. (2004). To test whether this result holds in the extremes of the performance measure properties, I use the estimates from Model 2_{TSLS} to determine the partial derivative of *DELEGATION* to *REL_INC_FPM* (see footnote 4):

$$\frac{\partial DELEGATION}{\partial REL\_INC\_FPM} \approx 0.05 + 0.20 \cdot REL\_SEN\_FPM + 0.19 \cdot REL\_PREC\_FPM + 0.00 \cdot REL\_VERIF\_FPM.$$

This partial derivative reflects the impact of *REL_INC_FPM* on *DELEGATION* as a function of the performance measure properties. If financial performance measures are poor incentive measures, i.e., at the minimum observed values of *REL_SEN_FPM* (-1.63) and *REL_PREC_FPM* (-2.31), delegation and the relative incentive use of financial performance measures are significantly negatively associated (slope ≈ -0.72; $p < 0.10$, two-tailed). In this setting, it is costly to contract on financial performance measures relative to nonfinancial performance measures and doing so creates an incentive-related cost of delegation, which leads to less delegation (cf., Nagar 2002). In contrast, if financial performance measures are good incentive measures, i.e., at the maximum observed values of *REL_SEN_FPM* (1.94) and *REL_PREC_FPM* (1.65), then delegation and the relative incentive use are significantly positively associated (slope ≈ 0.75; $p < 0.10$, two-tailed).[21] In this setting, financial performance measures complement the delegation choice, which leads to more delegation (cf., Abernethy et al. 2004).

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[19] Given that Equation (7) examines interactions between an endogenous variable and multiple exogenous variables, the "standard" application of the Durbin-Wu-Hausman test (e.g., using the first-stage residuals in the second stage) cannot be applied. I therefore return to the basics of the test, i.e., test whether there is a significant difference between the OLS estimator and TSLS estimator, by using a bootstrapping procedure. More specifically, I use 2,000 iterations with random resampling and calculate, for each sample, the difference between the OLS estimator and the HV-estimator and compute bias-corrected confidence intervals, which is used to determine the statistical significance of the difference.

[20] The finding that there are significant interaction effects in Model 2 essentially implies that Model 1 is misspecified. Although it is difficult to predict the consequences of this misspecification, it might explain why the relative precision of financial performance measures is significantly negatively related to delegation in Model 1, while such an effect is absent in Model 2.

[21] The difference between the negative slope (-0.72) and the positive slope (0.75) is statistically significant at the 5 percent level (two-tailed).

These results provide further support for the prediction that firms design their decision-making process around the quality of contractible performance measures (Prendergast 2002; Holmstrom and Milgrom 1994; Grossman and Hart 1986). When financial performance measures are able to solve the incentive problem, firms can delegate decision-making authority; but when it is difficult or costly to contract on financial performance, firms reduce the need for incentive contracting by lowering the extent of delegation. In addition, these results are important as they provide a plausible explanation for why previous research has been unable to find an effect of incentives on delegation. For example, Nagar (2002) finds that incentive-based pay, which, in his setting, is predominantly based on earnings, does not affect delegation. Similarly, Abernethy et al. (2004) and Demers et al. (2004) find that the relative use of financial performance measures for incentive purposes does not affect the extent of delegation. All three papers ignore the impact of performance measure properties and can, therefore, be considered an "on average" analysis (from a performance measure properties perspective). The results in this paper suggest that the average effect is likely to be zero because there are countervailing forces at play, which are driven by the performance measure properties.

Overall, the results are consistent with the hypothesis that the association between delegation and the relative use of financial performance measures for incentive purposes is positively affected by the relative quality of these measures.[22]

Performance Measure Properties and Incentive Use

To provide more insight into why sensitivity and precision have an interactive effect on delegation while verifiability has a direct effect, I examine the impact of the performance measure properties on the relative incentive use of financial performance measures, after controlling for the two dummy variables that represent the functional background of the manager and (the endogenous choice of) delegation. I use the variables related to the marginal benefits of delegation, described in Section IV, as instruments for the delegation choice. More specifically, I use *ENV_UNC_IN*, *ENV_UNC_OUT*, *SIZE*, and *HIER_LEVELS* as instruments. I do not expect these variables to have an impact on the relative incentive use of financial performance measures over and above the impact of delegation and specification tests confirm this expectation.[23]

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[22] A limitation underlying this conclusion is that I treat the performance measure properties as exogenous in the empirical model. However, it is possible that firms actually design the performance measure properties conditional on the decision to delegate, i.e., delegation affects the properties. If so, then this suggests a system of (three) equations in which the choice variables are simultaneously determined. Ignoring the endogeneity of the performance measure properties then potentially causes a simultaneity bias. The sign and magnitude of this bias is generally unknown *ex ante*, since it requires knowledge of the (true) parameters one is trying to estimate. Unfortunately, given the lack of valid instruments, I am econometrically unable to take the endogeneity of the performance measure properties into account. However, the core of the problem is that delegation potentially affects the performance measure properties. To examine this, I regress each performance measure property, i.e., *REL_SEN_FPM*, *REL_PREC_FPM*, and *REL_VERIF_FPM*, on the firm dummies and *DELEGATION* using TSLS, with *ENV_UNC_IN*, *ENV_UNC_OUT*, *SIZE*, and *HIER_LEVELS* as instruments. The results (not tabulated) show that the extent of delegation does not affect any of the three relative properties of financial performance measures. Although these results do not rule out the potential endogeneity of the performance measure properties, they do suggest that it is highly unlikely that the results presented in this paper are driven by reversed causality and/or simultaneity.

[23] The test of over-identifying restrictions indicates that the null hypothesis of exogenous instruments cannot be rejected ($p = 0.571$). In addition, the test of the first-stage exclusion restriction is significant ($p < 0.01$) and the partial R^2 equals 14 percent, which provides evidence of relevance.

Table 6 presents the results of the TSLS analysis, which indicates that *REL_INC_FPM* is positively affected by *REL_SEN_FPM* and *REL_PREC_FPM*, but not affected by *REL_VERIF_FPM*.[24] That is, the more sensitive and precise financial performance measures are relative to nonfinancial performance measures, the greater their relative use for incentive purposes, which is consistent with the analytical literature (e.g., Banker and Datar 1989) and previous empirical evidence (e.g., Abernethy et al. 2004). The finding that sensitivity and precision affect the incentive use of financial performance measures, while verifiability does not, provides a plausible explanation for why the former properties and the incentive use of financial performance measures positively reinforce each other in determining the extent of delegation, while this reinforcing effect is absent regarding verifiability.[25]

Table 6 further shows that *DELEGATION* has a significant positive effect on *REL_INC_FPM*, which is consistent with the expectation that delegation increases the relative demand for more aggregate performance measures (e.g., Abernethy et al. 2004). Finally, the dummy variable *SALES_MAN* has a significant positive effect on *REL_INC_FPM*, as expected, but the dummy variable *R&D_MAN* has no significant effect.

Aggregation versus Performance Measure Properties

The explanation that I provide for the positive interaction effects in Table 5 assumes that it is specifically the aggregate nature of financial performance measures that can complement the delegation choice. However, an alternative explanation for the interaction effects is that the incentive properties matter, irrespective of the type of performance measure and its level of aggregation. That is, the relative incentive use of any of the three types of performance measures increases delegation if these measures have relatively good incentive properties. If this alternative explanation holds, then it should not matter whether we construct the relative incentive use and properties as (1) x relative to y and z, (2) y relative to x and z, or (3) z relative to x and y. To examine the alternative explanation, I therefore re-estimate Equation (7) twice, and replace the relative incentive use and properties of financial performance measures (x versus y and z) by:

(1) the incentive use and properties of internal nonfinancial performance measures relative to those of financial and external nonfinancial measures (y versus x and z);
(2) the incentive use and properties of external nonfinancial performance measures relative to those of financial and internal nonfinancial measures (z versus x and y).

If the alternative explanation holds and only the incentive properties matter, then similar positive interaction effects should be observed after re-estimating Equation (7). If, however, it is the aggregate nature of financial performance measures versus the specific nature of nonfinancial performance measures that matters, in addition to the incentive properties, then no interaction effects or potentially negative interaction effects should be found.

The results (not tabulated) indicate that, for the internal nonfinancial performance measures, there are no significant interaction effects (one positive, two negative) between the relative incentive use and relative incentive properties on delegation. For the external nonfinancial performance measures, the interactions between the relative incentive use and all

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[24] Appendix B presents the results of the first-stage reduced form equation for delegation.

[25] The observation that, of all three performance measure properties, *REL_SEN_FPM* is the most significant determinant of *REL_INC_FPM* might explain why the OLS coefficient for the interaction term *REL_INC_FPM* × *REL_SEN_FPM* in Model 2_{OLS} has a significant endogeneity bias (see Table 5).

TABLE 6
Two-Stage Least-Squares Estimation of the Impact of the Extent of Delegation and Properties of Financial Performance Measures on the Relative Incentive Use of Financial Performance Measures
(n = 105)

Independent Variables	***REL_INC_FPM***
Intercept	−0.32*
$Firm_A$	−0.02
$Firm_B$	0.62**
$Firm_C$	0.18
$Firm_D$	−0.13
$Firm_E$	0.86*
DELEGATION#	0.48**
REL_SEN_FPM	0.47***
REL_PREC_FPM	0.09†
REL_VERIF_FPM	0.01
SALES_MAN	0.52***
R&D_MAN	0.10
Adjusted R^2	0.38***
Incremental R^2	0.18***
Partial R^2	0.24***

***, **, *, †Significant at the 1 percent, 5 percent, 10 percent, and 15 percent levels, respectively (two-tailed test using bias-corrected bootstrapped confidence intervals).

#Indicates significant endogeneity at the 5 percent level based on the Durbin-Wu-Hausman test.

The incremental R^2 indicates the total variance explained by the independent variables incremental to firm effects. The partial R^2 indicates how much of the within-firm variance is explained by the independent variables. The instrumental variables for the endogenous variable *DELEGATION* are *ENV_UNC_IN*, *ENV_UNC_OUT*, *SIZE*, and *HIER_LEVELS*. The first-stage adjusted R^2 equals 33 percent and the partial R^2 equals 14 percent. The test of the first-stage exclusion restriction is significant ($p < 0.01$), while the test of over-identifying restrictions indicates that the null-hypothesis of exogenous instruments cannot be rejected ($p = 0.571$). This suggests that the instruments are relevant and exogenous.

See Appendix A for variable measurement and definitions.

three relative incentive properties are negative, two of which are significant (sensitivity and verifiability). These findings are consistent with the theory proposed in Section II and inconsistent with the alternative explanation.

As a final test, I re-estimate Equation (7) and replace the relative incentive use and properties of financial performance measures (x versus y and z) by the incentive use and properties of external nonfinancial performance measures relative to those of internal nonfinancial performance measures (z versus y), i.e., one type of (specific) nonfinancial measure versus another type of (specific) nonfinancial measure. In line with the above results, I find no significant interactions effects on delegation. Overall, these results lend support for the argument that, in addition to the incentive properties, it is the aggregate nature of financial performance measures that allows for delegation.

VI. SUMMARY AND CONCLUSION

In this paper, I examine the determinants of delegation. I find that the contractibility of financial performance measures *vis-à-vis* nonfinancial performance measures increases delegation. More specifically, the empirical results show that, if financial performance measures are good incentive measures, i.e., have relatively high sensitivity, precision, and verifiability, then using these measures for incentive purposes can complement the delegation choice, which results in increased delegation. Furthermore, I find that delegation is affected by variables related to the internal and external environment, which is consistent with the argument that delegation increases when the marginal benefits of delegation are higher.

The results of this study are subject to several caveats. First, the variables used in the empirical analysis are based on the managers' perceptions. There is a possibility that the managers' perceptions of the variables of interest to this study differ from their superiors' perceptions. Although I use proprietary archival data to validate the survey-based measure of the incentive use of performance measures, which reduces the likelihood of these discrepancies, I cannot rule out that differences in perceptions exist. However, the likelihood that potential misperceptions drive the results is low given that these misperceptions should also be systematic. Second, I only examine a subset of organizational design variables and the analysis is therefore a partial equilibrium analysis (cf., Nagar 2002). Third, in any empirical analysis of choice variables on both the right-hand side and left-hand side of the equation, endogeneity remains an issue. Although I try to minimize the potential problems associated with endogeneity, by carefully selecting instruments, testing the validity of these instruments, and performing additional analyses, I cannot rule out the possibility that endogeneity is present and the model misspecified. Fourth, all of the firms that participate in this study are, or once were, clients of Hay Management Consultants. To the extent that these firms have incentive systems that differ from incentive systems of "non-clients," the results might not be generalizable to firms that have not been "assisted" by compensation consultants. Fifth, in the theoretical and empirical analysis of this paper, I assume that incentive contracting is the only means to address the delegation-incentive problem. Other common mechanisms exist such as the budgeting and resource allocation process that might also be able to address this problem. Future research can examine to what extent these other mechanisms play a role in the delegation-incentive problem. Finally, even though financial measures are more aggregate than nonfinancial measures, financial measures themselves can be more aggregate or less aggregate (more specific). Furthermore, these measures may vary along the dimensions of sensitivity, precision, and verifiability due to, for example, cost allocations and asset allocations. Future research can examine to what extent the incentive use and properties of more aggregate financial measures relative to less aggregate *financial* measures affect the delegation choice.

Despite the limitations, this study makes an important contribution to the accounting literature. Consistent with economic theory, it shows that firms design their decision-making process around the quality of contractible performance measures, and particularly around the contractibility of financial performance measures. It further provides evidence that the organizational design variables are simultaneously determined, which is consistent with arguments made in the theoretical literature (e.g., Brickley et al. 1997; Baiman and Rajan 1995; Melumad et al. 1992; Milgrom and Roberts 1992).

APPENDIX A
Variable Definitions and Measurement Instruments

DELEGATION = the extent to which decision-making authority is delegated to lower-level managers (measured by items a-e below).

Items	**Answering Format (1–5)**
Who has the decision-making authority with respect to ...	
a. Development of new products	superior's decision–my decision
b. Hiring and firing of personnel	superior's decision–my decision
c. Selection of large investments	superior's decision–my decision
d. Budget allocations	superior's decision–my decision
e. Pricing decisions	superior's decision–my decision

REL_INC_FPM = the relative use of financial performance measures for incentive purposes (measured by the difference between financial and nonfinancial measures for items a–h below).

Items	**Answering Format (1–5)**
How much importance does your superior attach to **xxx** performance measures in ...	
a. The evaluation of your performance	no importance–very high importance
b. Periodic performance reports	no importance–very high importance
c. Officially rating your performance	no importance–very high importance
d. Periodic discussions with you	no importance–very high importance
e. Determining your salary increases	no importance–very high importance
f. Determining your annual bonus	no importance–very high importance
g. Increasing your chance of promotion	no importance–very high importance
h. Increasing your authority within the organization	no importance–very high importance

REL_SEN_FPM = the relative sensitivity of financial performance measures to managerial actions (measured by the difference between financial and nonfinancial measures for items f–k below);

REL_PREC_FPM = the relative precision of financial performance measures (measured by the difference between financial and nonfinancial measures for items a and c–e below); and

REL_VERIF_FPM = the relative verifiability (objectivity and verifiability) of financial performance measures (measured by the difference between financial and nonfinancial measures for items l–n below).

Items	**Answering Format (1–5)**
My performance expressed in **xxx** performance measures is strongly affected by:	
a. ... changes in economic conditions	fully disagree–fully agree
b. ... decisions made in other parts of the organization	fully disagree–fully agree
c. ... changes in the behavior of customers	fully disagree–fully agree
d. ... changes in the behavior or strategies of suppliers	fully disagree–fully agree
e. ... changes in the behavior or strategies of competitors	fully disagree–fully agree
f. Whether I function well or not as a manager can be expressed accurately in **xxx** performance	fully disagree–fully agree

g. Many of the activities and tasks that I perform do not show up in **xxx** performance	fully disagree–fully agree
h. If I perform well as a manager, it is directly reflected in better **xxx** performance	fully disagree–fully agree
i. Working hard leads to better **xxx** performance	fully disagree–fully agree
j. Devotion and effort in my job leads to better **xxx** performance	fully disagree–fully agree
k. Providing effort in my job leads to better **xxx** performance	fully disagree–fully agree
l. The measurement of **xxx** performance is objective and verifiable	fully disagree–fully agree
m. The measurement of **xxx** performance is done by objective persons	fully disagree–fully agree
n. An independent person verifies the measurement of **xxx** performance	fully disagree–fully agree
o. The measurement of **xxx** performance is predominantly of a quantitative nature	fully disagree–fully agree

xxx reflects "financial," "external nonfinancial," or "internal nonfinancial."

ENV_UNC_IN = environmental uncertainty related to the input-side of the firm (measured by items c and d below); and
ENV_UNC_OUT = environmental uncertainty related to the output-side of the firm (measured by items a and b below).

Items	**Answering Format (0 [N.R.]; 1–5)**
I. Impact of ... on your job and unit	
a. Behavior and/or buying patterns of customers	no impact–very high impact
b. Behavior and/or strategies of competitors	no impact–very high impact
c. Technological developments in your profession	no impact–very high impact
d. Behavior and/or strategies of your suppliers	no impact–very high impact
e. Legal and/or political developments	no impact–very high impact
II. Predictability of (changes in) ...	
a. Behavior and/or buying patterns of customers	very predictable–very unpredictable
b. Behavior and/or strategies of competitors	very predictable–very unpredictable
c. Technological developments in your profession	very predictable–very unpredictable
d. Behavior and/or strategies of your suppliers	very predictable–very unpredictable
e. Legal and/or political developments	very predictable–very unpredictable

SIZE = ratio of the number of employees within the manager's unit to the total number of employees; and
HIER_LEVELS = dummy variable that equals 1 if there are multiple hierarchical levels below the manager's unit, and 0 otherwise.

APPENDIX B

TABLE B1
Reduced Form Equations for the Relative Incentive Use of Financial Performance Measures and the Extent of Delegation
(n = 105)

Independent Variables	*REL_INC_FPM*	*DELEGATION*
Intercept	−0.64***	−0.60***
$Firm_A$	0.35†	0.92***
$Firm_B$	0.87***	0.55***
$Firm_C$	0.53**	0.63***
$Firm_D$	−0.09	0.23
$Firm_E$	1.05**	0.37
ENV_UNC_IN	−0.06	0.10*
ENV_UNC_OUT	0.01	−0.01
SIZE	0.74	1.92**
HIER_LEVELS	0.29*	0.39**
REL_SEN_FPM	0.41***	−0.11*
REL_PREC_FPM	0.05	−0.11*
REL_VERIF_FPM	0.07	0.16**
SALES_MAN	0.56***	−0.02
R&D_MAN	0.18	−0.00
Adjusted R^2	0.44***	0.33***
Partial R^2	0.10***	0.14***

***, **, *, † Significant at the 1 percent, 5 percent, 10 percent, and 15 percent levels, respectively (two-tailed test using bias-corrected bootstrapped confidence intervals).
See Appendix A for variable measurement and definitions.

REFERENCES

Abernethy, M. A., J. Bouwens, and L. Van Lent. 2004. Determinants of control system design in divisionalized firms. *The Accounting Review* 79: 545–570.

Arya, A., J. Glover, and P. Liang. 2004. Intertemporal aggregation and incentives. *European Accounting Review* 13: 643–657.

Baiman, S., D. F. Larcker, and M. V. Rajan. 1995. Organizational design for business units. *Journal of Accounting Research* 33: 205–230.

———, and M. V. Rajan. 1995. Centralization, delegation, and shared responsibility in the assignment of capital investment decision rights. *Journal of Accounting Research* 33: 135–164.

Baker, G. P., and T. N. Hubbard. 2003. Make vs. buy in trucking: asset ownership, job design, and information. *American Economic Review* 93: 551–572.

Banker, R. D., and S. M. Datar. 1989. Sensitivity, precision, and linear aggregation of signals for performance evaluation. *Journal of Accounting Research* 27: 21–39.

Bound, J., D. A. Jaeger, and R. M. Baker. 1995. Problems with instrumental variables estimation when the correlation between the instruments and the endogenous explanatory variable is weak. *Journal of the American Statistical Association* 90: 443–450.

Brickley, J. A., C. W. Smith, and J. L. Zimmerman. 1997. *Managerial Economics and Organizational Architecture*. Burr Ridge, IL: Irwin.

Bushman, R. M., R. J. Indjejikian, and A. Smith. 1995. Aggregate performance measures in business unit manager compensation: The role of intra-firm interdependencies. *Journal of Accounting Research* 33: 101–134.

Christie, A. A., M. P. Joye, and R. L. Watts. 2003. Decentralization of the firm: Theory and evidence. *Journal of Corporate Finance* 9: 3–36.

Demers, E. A., M. B. Shackell, and S. K. Widener. 2004. Incentives, performance measurement, and delegation: An empirical analysis of interdependencies in organizational architecture. Working paper, University of Rochester, University of Notre Dame, and Rice University.

Dillman, D. A. 1978. *Mail and Telephone Surveys: The Total Design Method.* New York, NY: John Wiley & Sons.

Feltham, G. A., and J. Xie. 1994. Performance measure congruity and diversity in multi-task principal/agent relations. *The Accounting Review* 69: 429–453.

Fisher, J. 1995. Use of nonfinancial performance measures. In *Readings in Management Accounting*, edited by S. M. Young, 329–335. Englewood Cliffs, NJ: Prentice Hall.

Flannery, T. P., D. A. Hofrichter, and P. E. Platten. 1996. *People, Performance, and Pay: Dynamic Compensation for Changing Organizations.* New York, NY: The Free Press.

Gordon, L. A., and V. K. Narayanan. 1984. Management accounting systems, perceived environmental uncertainty and organizational structure: An empirical investigation. *Accounting, Organizations and Society* 9: 33–47.

Govindarajan, V. 1984. Appropriateness of accounting data in performance evaluation: An empirical investigation of environmental uncertainty as an intervening variable. *Accounting, Organizations and Society* 9: 125–135.

Grice, J. W., and R. J. Harris. 1998. A comparison of regression and loading weights for the computation of factor scores. *Multivariate Behavioral Research* 33: 221–247.

Grossman, S. J., and O. D. Hart. 1986. The costs and benefits of ownership: A theory of vertical and lateral integration. *Journal of Political Economy* 94: 691–719.

Hartmann, F. G. H., and F. Moers. 1999. Testing contingency hypotheses in budgetary research: An evaluation of the use of moderated regression analysis. *Accounting, Organizations and Society* 24: 291–315.

Heckman, J., and E. Vytlacil. 1998. Instrumental variables methods for the correlated random coefficient model. *Journal of Human Resources* 33: 974–987.

Holmstrom, B. 1979. Moral hazard and observability. *Bell Journal of Economics* 10: 74–91.

———, and P. Milgrom. 1987. Aggregation and linearity in the provision of intertemporal incentives. *Econometrica* 55: 303–328.

———, and ———. 1994. The firm as an incentive system. *American Economic Review* 84: 972–991.

———. 1999. The firm as a subeconomy. *Journal of Law, Economics and Organization* 15: 74–102.

Hubbard, T. N. 2000. The demand for monitoring technologies: The case of trucking. *Quarterly Journal of Economics* 115: 533–560.

Ittner, C. D., and D. F. Larcker. 1998. Innovations in performance measurement: Trends and research implications. *Journal of Management Accounting Research* 10: 205–238.

———, ———, and M. V. Rajan. 1997. The choice of performance measures in annual bonus contracts. *The Accounting Review* 72: 231–255.

———, and ———. 2002. Determinants of performance measure choices in worker incentive plans. *Journal of Labor Economics* 20: S58–S90.

Jensen, M. C., and W. Meckling. 1992. Specific and general knowledge structure. In *Contract Economics*, edited by L. Werin, and H. Wijkander, 251–274. Oxford, U.K.: Basil Blackwell Press.

———. 2001. Value maximization, stakeholder theory, and the corporate objective function. *Journal of Applied Corporate Finance* 14: 8–21.

Jeong, J., and G. S. Maddala. 1993. A perspective on application of bootstrap methods in econometrics. In *Handbook of Statistics*, Vol. 11, edited by G. S. Maddala, C. R. Rao, and H. D. Vinod, 573–610. Amsterdam, The Netherlands: Elsevier Science.

Keating, A. S. 1997. Determinants of divisional performance evaluation practices. *Journal of Accounting and Economics* 24: 243–273.

Khandwalla, P. N. 1972. The effect of different types of competition on the use of management controls. *Journal of Accounting Research* 10: 275–285.

Larcker, D. F. 1981. The perceived importance of selected information characteristics for strategic capital budgeting decisions. *The Accounting Review* 56: 519–538.

———, and T. O. Rusticus. 2004. On the use of instrumental variables in accounting research. Working paper, The University of Pennsylvania.

Libby, T., and J. H. Waterhouse. 1996. Predicting change in management accounting systems. *Journal of Management Accounting Research* 8: 137–150.

Melumad, N. D., D. Mookherjee, and S. Reichelstein. 1992. A theory of responsibility centers. *Journal of Accounting and Economics* 15: 445–484.

———, ———, and ———. 1995. Hierarchical decentralization of incentive contracts. *RAND Journal of Economics* 26: 654–672.

Merchant, K. A. 1990. The effects of financial controls on data manipulation and management myopia. *Accounting, Organizations and Society* 15: 297–313.

Milgrom, P., and J. Roberts. 1992. *Economics, Organization and Management.* Englewood Cliffs, NJ: Prentice Hall.

Nagar, V. 2002. Delegation and incentive compensation. *The Accounting Review* 77: 379–395.

Prendergast, C. 2002. The tenuous trade-off between risk and incentives. *Journal of Political Economy* 110: 1071–1102.

Ramsey, J. B. 1969. Tests for specification errors in classical linear least-squares analysis. *Journal of the Royal Statistical Association B* 71: 350–371.

Williamson, O. E. 1975. *Markets and Hierarchies: Analysis and Antitrust Implications.* New York, NY: The Free Press.

Wooldridge, J. M. 2002. *Econometric Analysis of Cross Section and Panel Data.* Cambridge, MA: MIT Press.

———. 2003. Further results on instrumental variables estimation of average treatment effects in the correlated random coefficient model. *Economics Letters* 79: 185–191.

Wruck, K. H., and M. C. Jensen. 1994. Science, specific knowledge, and total quality management. *Journal of Accounting and Economics* 18: 247–287.

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The Perils of Pensions: Does Pension Accounting Lead Investors and Analysts Astray?

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ABSTRACT: This paper explores the ability of investors and analysts to fully process available pension information when establishing prices and making earnings forecasts. I find that neither prices nor forecasts fully reflect the quantifiable future earnings effects of changes in pension information at the time it becomes publicly available in the firm's 10-K. Instead, the evidence suggests that investors and analysts only gradually incorporate this information into prices and forecasts as they observe the effects of the pension plan changes on subsequent quarterly earnings. The persistent tardiness of analysts to incorporate this relevant and economically significant information about earnings is surprising given that they are provided with pension information on a repeated and timely basis. Additionally, I find that the off-balance-sheet portion of the pension plan's funded status and the PBO are predictive of future returns while the on-balance-sheet portion of the funded status is not. This implies that investors do not accurately assess the long-run cash flow and earnings implications of these off-balance-sheet pension disclosures. The compensation rate is also predictive of future returns suggesting that it signals management's expectation of future performance.

Keywords: *pensions; analysts; forecasts; earnings management.*

Data Availability: *Data are publicly available from sources identified in the paper.*

I. INTRODUCTION

This paper explores whether investors and analysts fail to fully process the disclosed pension information available to them under SFAS No. 87 (FASB 1985) when establishing prices and making earnings forecasts. My primary goal is to identify

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specific pension accounting information that analysts and investors fail to assimilate completely. Identification of this information can help market participants better evaluate pension disclosures, highlight areas of pension accounting that firms may utilize to manage earnings, and serve as a benchmark for standard setters to better assess the usefulness of future required pension disclosure. This is particularly relevant as pension expense (or income) has become a major component of net income for firms with defined benefit pension programs and because a re-evaluation of how best to account for pension plans is one of the next issues to be addressed by the Financial Accounting Standards Board (FASB) and the International Accounting Standards Board (IASB) in the movement toward international accounting convergence. This paper also aids in the general understanding of how analysts generate forecasts and, in particular, whether they *repeatedly* fail to fully incorporate accounting information, even when its future earnings impact can be reliably estimated.

Past research (Plumlee 2003; Hirst and Hopkins 1998) has shown that information complexity and presentation can affect both the assimilation and weighting of information by analysts. Because pension accounting is technically complex and because most pension information is disclosed in involved footnotes rather than reported on the income statement or balance sheet, one might expect both analysts and investors to have difficulty assimilating relevant information. Aside from the complexity of pension accounting, investors have also expressed concern that the amount of information disclosed is inadequate to fully gauge the impact of defined benefit pension plans on future firm performance. This concern prompted the Financial Accounting Standards Board (FASB) to revise SFAS No. 132 in December 2003 with the objective of helping investors to better understand the risks and cash flow implications of defined benefit pension plans.[1]

The SFAS No. 132 revision specifically addresses problems of information availability by increasing the frequency (quarterly instead of annually) as well as the scope (e.g., tabular form for reporting pension assumptions, breakdown of plan assets by asset class, disclosure of the accumulated benefit obligation [ABO], and a payout schedule for future benefits) of pension disclosures. While this new information will enable investors and analysts to better asses the impact of a firm's pension plan on firm performance, it will only be effective to the extent that it can be effectively processed by investors. By detailing the extent to which both analysts and investors incorporate previously available pension information, this paper provides a benchmark that will aid in evaluating the impact of the new revision.

Prior research (e.g., Bernard and Thomas 1990; Abarbanell and Bernard 1992; Elgers and Lo 1994; Abarbanell and Bushee 1997, 1998; Bradshaw et al. 2001) has demonstrated that analysts fail to completely incorporate empirically observed relationships between accounting information and earnings when generating forecasts. However, as the cause of these empirical "anomalies" and their expected persistence is not yet fully understood, it is unclear to what extent analysts would be expected to incorporate these observed relationships into forecasts, especially on a per-firm basis. What is perhaps more surprising are the findings of Biddle and Ricks (1988), Chen and Schoderbek (2000), and Plumlee (2003) in which analysts fail to fully adjust forecasts to reflect the effects of accounting and tax law changes whose impact on an individual firm can be reliably estimated. While these results demonstrate that analysts may have trouble initially assessing the impact of specific *one-time* accounting changes, they do not demonstrate that analysts have trouble assessing the impact of accounting information that they deal with on a *recurring* basis. This paper thus extends this line of literature by demonstrating that analysts do not fully reflect the recurring and reliably predictable effects of changes in pension information. These findings provide insight into analyst sophistication and how well forecasts reflect more subtle and complex accounting information.

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[1] The FASB motivates the new revision as follows: "Although current disclosure requirements for pension plans and other postretirement benefit plans are extensive, many users of financial statements told the Board that the information provided for defined benefit pension plans is not adequate. Users of financial statements requested additional information that would assist them in (a) evaluating plan assets and the expected long-term rate of return used in determining net pension cost, (b) evaluating the employer's obligations under pension plans and the effects of such obligations on the employer's future cash flows, and (c) estimating the potential impact of net pension cost on future net income. The Board concluded that disclosures about pensions could be improved to provide information that would better serve the users' needs" (FASB Exposure Draft 1025-200, FASB 2003, 1).

The first part of this paper investigates whether analysts and investors readily process *changes* in pension plan parameters that have a quantifiable effect on future earnings. I consider changes in the pension plan rate assumptions, the projected benefit obligation (PBO), and the pension plan assets (PA). These items, disclosed in a firm's 10-K, have a mechanical effect on one-year-ahead earnings that can be estimated at least nine months prior to the actual earnings announcement. I examine whether analysts expeditiously incorporate these expected earnings changes into their year-ahead earnings forecasts, and whether these changes are predictive of quarterly forecast errors.[2] I also explore whether these pension changes are predictive of stock returns over the next 12 months.

I find that analysts do not explicitly incorporate pension parameter change information into their initial forecasts, and that these changes are predictive of quarterly earnings forecast errors in the following three quarters. My findings show that analyst forecasts only gradually reflect this pension information over the next eight months. Analysts thus appear to incorporate this pension information by observing its effect on subsequent quarterly earnings announcements rather than utilizing the pension notes to directly estimate the earnings effects. The failure of analysts to incorporate pension information appears to be driven by the complexity of the task rather than a lack of information, because all of the necessary information is readily available in the firm's 10-K.

Like analysts, the stock market also does not appear to fully incorporate pension change information at the time of release. I find that pension parameter changes are predictive of returns in the year following the information release, and that a hedge portfolio based on the changes yields approximately a 4.6 percent return. The magnitude and consistency of these returns demonstrate that the mechanical earnings effects of pension plan changes are of sufficient importance to be worth the attention of both analysts and investors.

The second part of the paper examines whether the market has difficulty estimating the long-run earnings and cash flow implications of various pension plan parameters and whether the *magnitude* of these parameters are therefore predictive of firms' short-run and long-run returns. The particular parameters I consider are the two components of the firm's funded status (hereafter, on-balance-sheet and off-balance-sheet liabilities), the firm's PBO (the total estimated pension liability), and the level of the three disclosed pension rate assumptions. Of these parameters only the PBO, off-balance-sheet liabilities, and the compensation rate prove consistently statistically predictive of future returns in various time intervals over the one- to five-year horizon. Specifically, I find that investors fail to properly account for the size of a firm's pension liability and the importance of the unfunded liability that has yet to be recognized on the income statement. Because the on-balance-sheet liability is not predictive of future returns, it appears that investors are unable to fully assess the information buried in the pension footnotes but can accurately assess the information that has been reflected in earnings. Additionally, I find that firms with higher compensation rate assumptions tend to experience higher future returns. This is interesting because higher compensation rates will cause a firm to recognize higher overall pension costs and thus lower earnings, all else being equal. Higher compensation rates may, therefore, serve as a bellwether for positive management expectations. In sum, these results indicate that the market is unable to fully impound salient pension plan information currently available to investors.

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[2] Forecast error is defined as actual earnings minus the corresponding analyst consensus forecast. Forecast errors at the time actual earnings are announced can also be interpreted as earnings surprises.

My examination makes two main contributions to existing knowledge. First, I document specific publicly available pension information that analysts and investors fail to fully incorporate into both forecasts and firm valuation. This provides a benchmark by which to evaluate the effect of the recent revision to SFAS No. 132. Analysts and investors can also use this knowledge to better predict future firm performance. Second, I demonstrate that analyst forecasts do not fully reflect the more subtle and complex aspects of pension accounting even though they deal with this information on a repeated basis and the earnings effects are readily quantifiable for individual firms.

Section II provides a brief background on the current state of defined benefit pension plans. Section III outlines my research questions in the context of related research. Section IV lays out my sample selection procedures and provides descriptive statistics. Section V provides a basic overview of my estimation methodology. Section VI presents the results of my analyses. Finally, Section VII provides a brief summary and conclusion.

II. BACKGROUND

In the last few years, defined benefit pension plans have gained increased attention from investors, regulators, and politicians. Several factors have led to the recent notoriety of pension plans: The relatively large value of the liabilities and assets in question; the number of firms with severely underfunded pension plans; the increasing cost to both workers and the government as pension plans are terminated; a complex system of pension accounting that keeps much of a pension plan's financial health off a firm's balance sheet and income statement; and the fact that small changes in assumptions can result in large changes in pension plan value. This last factor makes the funded status of pension plans extremely volatile and pension accounting an area ripe for potential earnings management.

With $140 billion in unfunded liabilities in the S&P 500 alone at the end of 2005 and the deficit in the Pension Benefit Guarantee Corporation (PBGC) reaching an unprecedented $25.7 billion after absorbing 120 new plan terminations, action is being taken on various fronts to shore up pension plan funding, curb potential accounting abuses, and revise pension accounting standards. Congress is debating legislation to strengthen pension plan funding requirements and increase PBGC premiums. The SEC has begun investigations into the pension accounting assumptions of several prominent firms such as General Motors, Boeing, Delphi, and Ford. Additionally, the FASB has recently voted to perform a two-stage comprehensive revision of pension accounting standards. The first stage, expected to be implemented in 2006, may require firms to put the entire funded status of their pension plans on the balance sheet. The second stage, likely to be conducted in concert with the IASB and to take at least three years, will reconsider all elements of pension accounting. It is notable that the announcement of the upcoming revision comes shortly after the recent revision to SFAS No. 132, just implemented in 2003.

III. LITERATURE REVIEW AND HYPOTHESIS DEVELOPMENT

The first part of this paper explores whether analysts and investors fail to recognize the predictable effects that *changes* in certain pension parameters have on future earnings. By performing fairly straightforward mechanical calculations based on SFAS No. 87, analysts and investors should be able to closely approximate these earnings effects on a per-firm basis. Analysts not recognizing these effects will fail to properly update their earnings forecasts; this will result in analyst forecast errors and corresponding surprises in realized quarterly earnings. Investors not properly assimilating this information will likewise be surprised when it manifests in quarterly earnings, making the pension parameter changes predictive of future stock returns.

The second part of this paper examines whether investors are able to understand fully the relation between the *magnitude* of various pension parameters and long-run firm earnings and cash flows. Estimating these long-run effects is much more difficult than mechanically calculating the short-term earnings effects of pension changes. If investors fail to completely incorporate disclosed pension information in current prices, the level of various pension parameters will have long-run return predictability.

Changes in Pension Plan Parameters

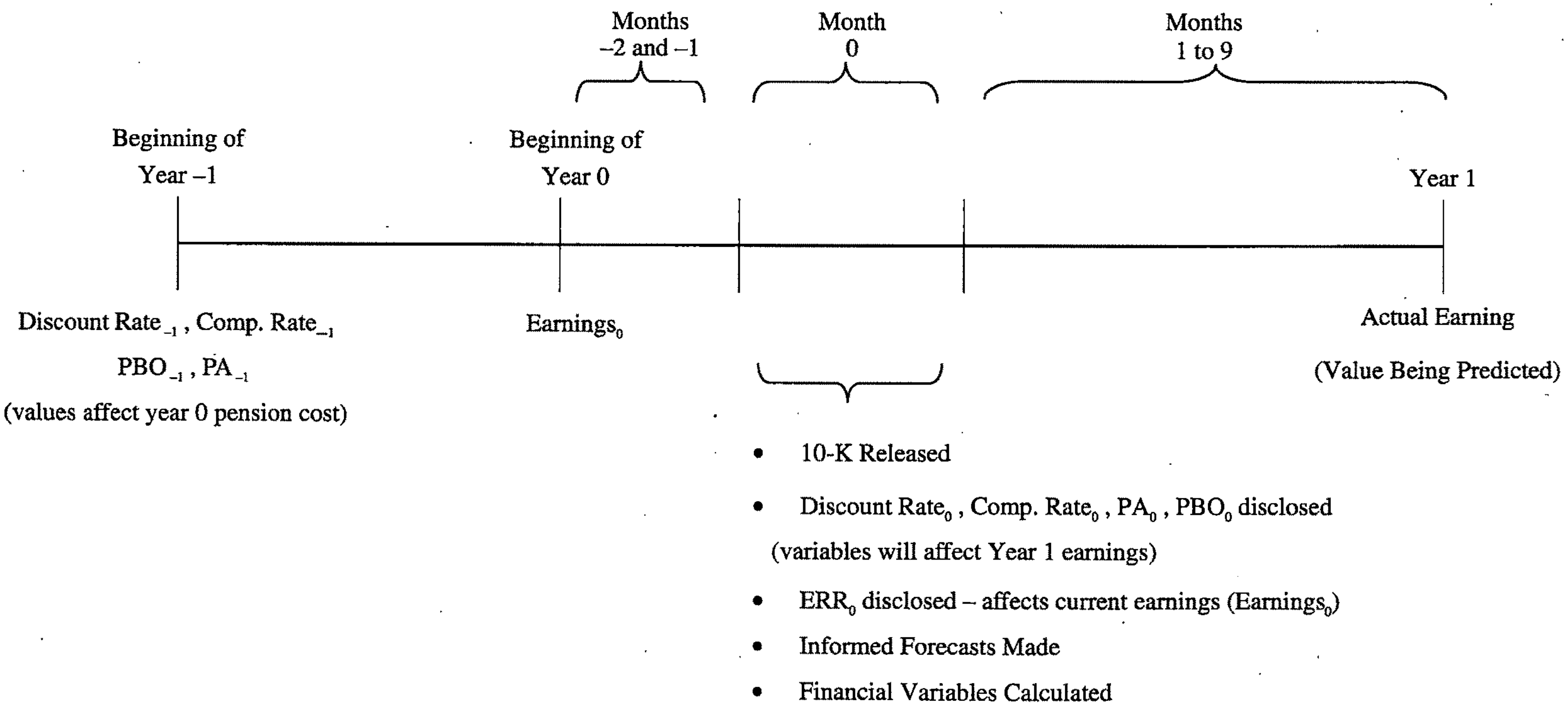
Changes in the PBO, plan assets, discount rate, and compensation rate have an impact on future pension cost, but do not affect concurrently reported earnings. This result occurs because SFAS No. 87 specifies that concurrent pension costs should generally be based on pension plan values reported in the previous year's 10-K. Changes in these pension plan parameters therefore have a mechanical predictive ability for future earnings beyond that of current earnings. If investors and analysts are aware of this relationship, then they should readily incorporate the new information into stock prices and forecasts, and the changes should not be predictive of either future earning surprises or returns. Figure 1 details the basic timeline when considering the effects of these pension plan changes.

The future pension cost components affected by changes in the various pension parameters are the service cost, the interest cost, and the expected return on plan assets. The service cost is affected by changes in both the compensation rate and the discount rate. Increases in the compensation rate increase the future expected payouts to employees and, thus, increase the yearly service cost. Increases in the discount rate cause future payouts to employees to be discounted at a higher rate and, thus, decrease the yearly service cost. The opposite effects occur when the rates are decreased.

The interest cost increases (decreases) when either the PBO or the discount rate increases (decreases). Because interest cost is essentially the increase in the PBO due to the passage of time, it can be approximated as the prior year's reported PBO times the prior year's discount rate. Similarly, the expected return on plan assets increases (decreases) when either the plan assets or the ERR increase (decrease). Unlike changes in the other two rates, however, changes in the ERR are reflected in the contemporary expected return on plan assets and, thus, in current earnings. Expected return on plan assets can, therefore, be approximated as the prior year's reported plan assets times the *current* year's ERR. This means that changes in a firm's plan assets will result in future earnings changes, while changes in ERR will be immediately reflected in earnings.

Using disclosed changes in a firm's discount rate, compensation rate, PBO, and plan assets, I calculate the expected change in the following year's service cost, interest cost, and expected return on plan assets. In order to gain additional insight into how the individual pension parameters affect future earnings, I divide the expected change in next year's interest cost into three components: the portion due to changes in the discount rate, the

FIGURE 1
Timeline for Estimating the Mechanical Earnings Effects of Pension Plan Changes


Months
−2 and −1
Month
0
Months
1 to 9
Beginning of
Year −1
Beginning of
Year 0
Year 1
Discount Rate$_{-1}$, Comp. Rate$_{-1}$
PBO$_{-1}$, PA$_{-1}$
(values affect year 0 pension cost)
Earnings$_0$
Actual Earning
(Value Being Predicted)
10-K Released
Discount Rate$_0$, Comp. Rate$_0$, PA$_0$, PBO$_0$ disclosed
(variables will affect Year 1 earnings)
ERR$_0$ disclosed – affects current earnings (Earnings$_0$)
Informed Forecasts Made
Financial Variables Calculated

portion due to the change in PBO, and a residual amount not directly related to either parameter. I also compute the total expected mechanical change in future earnings by summing all the individual cost components calculated. This value is referred to as total mechanical change in earnings (TOTMEC). The calculation of each cost component is explained in greater detail in Section V.

Mikhail et al. (1999) demonstrate that earnings forecast accuracy is important to analysts and that analysts become more accurate as they gain firm-specific experience. Analysts should therefore be motivated to utilize changes in pension information to increase their forecast accuracy provided the associated effort costs are not prohibitive. If analysts readily incorporate changes in pension information into their forecasts, then estimated changes in pension cost components should not be predictive of analyst forecast errors. However, if analysts only gradually incorporate this information by observing its effect on subsequent quarterly earnings, then these pension plan changes will predict forecast errors for both annual and quarterly earnings over the nine months following the release of a firm's 10-K. Hypotheses 1 and 2 (both stated in alternate form) test whether analyst forecasts only gradually reflect the earnings effects of changes in pension parameters:

H1: Changes in individual pension cost components are predictive of analyst forecast errors of annual earnings in the months after firms' 10-K release dates.

H2: The total expected mechanical change in future earnings due to changes in pension plan parameters (TOTMEC) will be predictive of future forecast errors of quarterly earnings.

Hypothesis 3 examines whether investors fully incorporate the future earnings effects of pension changes into stock prices. If investors account for these effects at the time the pension information is released, then the changes should not be predictive of future returns. If, on the other hand, investors do not fully process this information, then future returns should gradually reflect investors' reaction to these unexpected future earnings. I therefore test Hypothesis 3 (stated in alternative form):

H3: Portfolios formed on the total expected mechanical change in future earnings due to pension plan parameters (TOTMEC) will be predictive of one-year-ahead future abnormal returns.

Magnitude of Disclosed Pension Plan Parameters

The post-SFAS No. 87 empirical research shows mixed results on whether the market is able to correctly incorporate the current level of pension plan information when valuing firms. Barth (1991) concludes that the market incorporates PBO, though noisily, when valuing firms rather than simply relying on pension measures that are reported on the balance sheet. This shows that the market pays some attention to the pension information reported in the financial statement notes, but may have trouble properly weighting this information. Barth et al. (1992) test whether the market assigns different coefficients to the various components of reported pension cost based on their perceived permanence. They find that, in general, the market weights the various components differently, although not necessarily with the magnitude or sign predicted. In particular, the service cost component has an anomalous sign and is not statistically significant in explaining the current stock price. This shows that investors may have trouble properly weighting pension costs when assessing firm value.

Brown (2002) suggests that the market is able to effectively unwind aggressive managerial assumptions and, thus, fairly values pension liabilities and assets. His findings indicate that investors can properly weight the various pension parameters, as well as adjust them, for perceived earnings manipulation. In contrast, Gold (2000) hypothesizes that the smoothing mechanism in SFAS No. 87 separates the risk of pension assets from their reported return. He develops a model in which investors fail to adjust for plan asset risk and rely on reported pension costs rather than the disclosed plan assets and liabilities. Coronado and Sharpe (2003) adopt Gold's (2000) theoretical model and find that investors tend to incorrectly judge the permanence of pension earnings and thus treat pension earnings as core earnings. Their results show that this has caused investors to overvalue companies with defined benefit pension programs, although this overvaluation has not been significant until 2001. They attribute Brown's (2002) contrary findings to his failure to account for pension cost accruals.

Overall, the existing literature is inconclusive as to whether the market is able to completely incorporate pension information into firm valuation. One potential problem with all of the studies to date is that they rely on some form of a pricing model to determine the accuracy of current valuation. The results are thus subject to misspecifications in the models used. Rather than regressing firms' pension parameters on current price, this paper tests whether contemporary pension parameters are predictive of future returns. If pension parameters are predictive of future returns, then this indicates that the market does not fully appreciate the earnings and cash flow implications of reported pension information and is, therefore, surprised by these relationships when they surface in the future.

I examine whether six different pension parameters are predictive of future returns: the PBO, the firm's funded status (decomposed into two values, the on-balance-sheet and off-balance-sheet liabilities), the ERR, the discount rate, and the compensation rate. PBO measures the total present value of a firm's pension obligation. The larger the value, the more susceptible the firm is to market conditions that affect pension plan health. If investors do not consider this large liability when considering a firm's other liabilities, then firms with larger projected benefit obligations could experience negative future returns. To test whether investors underestimate price implications of the PBO, I examine the following hypothesis (stated in its alternative form):

H4: Partitioning firms based on their PBO scaled by total assets will result in negative abnormal future returns.

Many practitioners view a pension plan's funded status as its most important feature because it indicates whether a plan has sufficient assets to cover its future expected payouts. Large shortfalls in funded status could signal serious reductions in a firm's future cash flows and a decrease in earnings as the shortfall is amortized. Likewise, plans with large surpluses may not require funding in the foreseeable future and the excess funding will boost earnings as it is amortized. If investors do not fully understand these earnings and cash flow considerations, then the more overfunded (underfunded) a firm is, the better (worse) the firm's long-run performance will be.

Hirst and Hopkins (1998) demonstrate that analysts are more likely to attend to information when it is clearly and distinctly presented. In particular, they show that information provided in nonprominent formats, such as the statement of changes in equity, is less likely to be incorporated into valuation and analysis. Similarly, Davis-Friday et al. (1999) provide evidence that post-retirement benefit (PRB) liabilities that are recognized receive greater weight than PRB liabilities that are disclosed. Based on this prior research, I partition each firm's funded status into two parts: an on-balance-sheet liability (or asset) that represents the size of the unfunded liability (or overfunded pension asset) that has been recognized in past earnings, and an off-balance-sheet liability (or asset) that represents the size of the unfunded liability (or overfunded pension asset) that has not yet rolled through a firm's income statement and is, therefore, not reflected on the balance sheet. I hypothesize that investors will correctly asses the value implications of the recognized portion of a firm's funded status (the on-balance-sheet liability), but fail to recognize the full value implications of the unrecognized portion (the off-balance-sheet liability). This results in the following two hypotheses, the first nondirectional, and the second directional and stated in its alternative form:

H5a: Partitioning firms based on their on-balance-sheet liability scaled by total assets will not yield future abnormal returns.

H5b: Partitioning firms based on their off-balance-sheet liability scaled by total assets will result in negative future abnormal returns.

The rate assumptions used by firms also have potential implications for future performance. Firms that use more aggressive rate assumptions will experience higher earnings than similar firms with more conservative rate assumptions. While in most earnings management situations aggressive assumptions soon reverse and result in lower future earnings, this is not necessarily the case for pension plan rate assumptions. If a firm is too aggressive in its ERR assumption, then the actual return on plan assets will fall below the expected return. This return difference will be accumulated in a pool of unrecognized gains and losses where they may be offset by other actuarial gains. Only if this unrecognized pool gets large enough (10 percent of plan assets or PBO, whichever is larger) will the amount over this 10 percent limit be slowly amortized into pension cost, generally over a period not less than ten years. The feedback loop to adjust for aggressive discount rate assumptions is similarly glacial. If investors do not account for the positive effects of aggressive discount rate and ERR assumptions, then aggressive firms will experience higher stock returns than similar firms with more conservative accounting practices and correspondingly lower earnings. If, however, investors become aware of the long-term consequences of inflated assumptions and discount appropriately, then aggressive firms may experience lower future returns. It is, thus, an empirical question as to whether higher ERR and discount rate assumptions will cause positive or negative abnormal returns in the future.

The possible long-run return consequences resulting from a firm adopting an aggressively low compensation rate are harder to predict than the discount rate and expected rate of return effects. On the one hand, the aggressive policy of choosing a lower compensation rate should boost earnings in a manner similar to the other two rates. On the other hand, because the compensation rate is a measure of expected salary growth, a higher reported compensation rate could signal management's optimistic appraisal of future performance. Because this positive signal could outweigh the earnings benefits of choosing an aggressively low compensation rate, it is unclear what the returns implication of a high compensation rates may be. I test the following nondirectional hypothesis:

H6: Partitioning firms based on the level of their ERR, discount rate, and compensation rate assumptions will not result in future abnormal returns.

IV. SAMPLE SELECTION AND DATA DESCRIPTION

The initial sample consists of all firms available on both the Compustat and CRSP databases and having a Compustat fiscal year from 1988 to 2001. I select 1988 as the beginning fiscal year because SFAS No. 87, the FASB statement that mandated many of the current pension reporting requirements, became effective for firms with fiscal years beginning after December 15, 1986. Beginning in 1988 ensures that all firms are reporting under SFAS No. 87 and complying with the new requirements. The data set ends in 2001 because this was the most current fiscal year at the time of the study containing the necessary long-run variables. I add analyst information from the I/B/E/S database, although firms are not eliminated if they do not have analyst information. I then eliminate all firms that have no earnings report dates listed in either Compustat or I/B/E/S. I also eliminate any firms not on the NYSE, AMEX, or NASDQ, all ADRs, firms with assets less than $10 million, firms with a year-end closing price below $2, firms with less than one year of returns data, and all utilities and banks (firms with four-digit SIC codes between 4900 and 5000 and between 6000 and 7000). The remaining data set consists of 51,451 firm-years and contains both firms with and without defined benefit pension plans. This expanded data set allows the control variable rankings used in regressions (such as book-to-market, total accruals, etc.) to be based on the larger universe of firms rather than just firms with defined benefit pension plans.

For the various tests in the paper, I use all observations that are not missing the required pension parameters and control variables. Table 1 presents some descriptive statistics for the firms with defined benefit pension plans. The first three columns show the number of

TABLE 1
Various Descriptive Statistics for Firms with Pension Plans

	Number of Firms in Compustat with				Average			
Year	Pension Plans	All Liabilities Reported	All Pension Rates Reported	Predicted Pension Cost Changes	PBO	On-Balance-Sheet Liabilities	Off-Balance-Sheet Liabilities	Total Estimated Change in Pension Cost (TOTMEC)
					(as a percent of total assets)			($ per share)
1988	1078	645	—	—	14.18	0.21	−1.87	—
1989	1051	813	—	—	14.57	0.17	−1.94	—
1990	1003	809	—	—	14.51	0.04	−0.67	—
1991	1049	821	421	—	20.39	1.59	−1.06	—
1992	1100	861	495	—	16.53	0.23	−0.51	—
1993	1149	901	544	518	18.24	0.002	0.48	0.018
1994	1202	947	628	523	16.77	−0.10	0.38	−0.045
1995	1205	952	619	526	17.97	−0.01	0.46	0.040
1996	1222	975	620	503	17.74	−0.08	−0.46	0.038
1997	1222	968	617	490	18.10	−0.30	−1.28	0.055
1998	1178	958	647	465	18.84	−0.43	−0.48	0.016
1999	1099	892	626	457	16.85	−0.52	−2.61	0.033
2000	1007	788	559	450	16.23	−0.78	−1.38	−0.043
2001	988	739	521	52	16.96	−0.94	2.06	−0.075
	15,553	12,069	6,297	3,984	16.99	−0.07	−0.63	0.004

firms each year with defined benefit pension plans (i.e., having a reported PBO), the number of firms for which I could calculate on- and off-balance-sheet liabilities, and those with all three pension rate variables available on Compustat. The fourth column shows the number of firms for which I am able to calculate estimated changes in future pension costs.[3] Table 1 also details the size of the various pension liabilities, as well as a yearly average of the total estimated future change in pension cost due to changes in the pension parameters (TOTMEC).

V. METHODOLOGY OF ESTIMATING FUTURE PENSION COST COMPONENTS

Future Interest Cost and Expected Return on Plan Assets Component Estimations

The basic methodology to estimate the future interest cost and expected return on plan assets components of pension cost is as follows:

Assume: (Figure 1 graphically displays this information)

$t = -1$ is the time of the previous year's 10-K release and earnings announcement.
$t = 0$ is the time of the current 10-K release and earnings announcement. It includes new PBO, plan asset, and rate assumptions (if these have changed).
$t = 1$ is the time of the following year's earnings announcement. These are the future earnings being predicted.

Accounting for pension plan changes reported in the current 10-K yields "informed" forecasts of the future interest cost and expected return on plan assets. These future pension cost components are estimated as follows:[4]

$$\text{Informed Interest Cost Forecast} = PBO_0 * DR_0 \tag{1}$$

$$\text{Informed Expected Return on PA Forecast} = PA_0 * ERR_0 \tag{2}$$

where:

DR = discount rate; and
ERR = expected rate of return on plan assets.

A "simple" forecast of these costs would assume that they maintain at the level just reported in the pension footnote (the actual interest cost and expected return on plan assets in year 0). The difference between the informed and simple forecasts can thus be expressed:

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[3] Compustat has the necessary information to compute the on- and off-balance-sheet liability and projected benefit obligation going back to 1988. Rate assumptions are not available on Compustat until 1991. Calculation of future changes in pension cost is not possible until 1993, two years after the rate assumptions are available.

[4] For my calculation of the expected return on plan assets, I use the firm's reported plan assets at time 0 (the only value reported in Compustat). Under SFAS No. 87, however, firms may also calculate this value using a weighted average of plan assets over the last five years. For firms using this smoothing mechanism, my estimates will be less precise and thus biased against yielding significant results. Davis-Friday et al. (2005) provide a good description of this smoothing process and how to estimate its effect on earnings with more detailed data.

$$\Delta \text{ Interest Cost} = \text{PBO}_0 * \text{DR}_0 - \text{Actual Interest Cost}_0 \quad (3)$$

$$\Delta \text{ Expected Return on PA} = \text{PA}_0 * \text{ERR}_0 - \text{Actual Expected Return}_0. \quad (4)$$

In order to determine what specific pension plan parameter changes are not fully incorporated by analysts, I express the actual interest cost and expected return on plan assets in terms of their expected value based on the current year's ERR and last year's (year −1's) discount rate, PBO and plan assets, plus a realized difference accounting for other pension plan activity such as plan contributions and payouts. In equation form, this representation is:

$$\text{Actual Interest Cost}_0 = \text{PBO}_{-1} * \text{DR}_{-1} + \text{Realized Difference} \quad (5)$$

$$\text{Actual Expected Return}_0 = \text{PA}_{-1} * \text{ERR}_0 + \text{Realized Difference}. \quad (6)$$

The realized difference can be calculated for the interest cost using available Compustat data. Calculating this value for expected return prior to 1998, however, is problematic. Before the adoption of SFAS No. 132 (FASB 1998), firms could report either their expected return on plan assets directly as a pension cost component, or their actual return on plan assets as a cost component and the difference between the actual and expected return as part of the unrecognized gains and losses component of pension cost. The Compustat data record only the "return" field on a firm's financial statement and does not specify whether this return is the actual or expected return. This makes the actual expected return value reported in Compustat unusable prior to 1998. As a result, I assume that the actual return on plan assets in any given year is the previous year's plan assets times the current ERR. This assumption therefore makes the realized difference in Equation (6) zero.

The interest cost forecast and expected return forecast (Equations (1) and (2)) can also be rewritten as functions of the previous year's PBO, plan assets and discount rate and the change in these values between year −1 and year 0 as follows:

$$\text{Informed Interest Cost Forecast} = (\text{PBO}_{-1} + \Delta\text{PBO}) * (\text{DR}_{-1} + \Delta\text{DR}) \quad (7)$$

$$\text{Expected Return on PA Forecast} = (\text{PA}_{-1} + \Delta\text{PA}) * \text{ERR}_0 \quad (8)$$

where:

ΔDR = change in discount rate $(\text{DR}_0 - \text{DR}_{-1})$;
ΔPBO = change in PBO $(\text{PBO}_0 - \text{PBO}_{-1})$; and
ΔPA = change in plan assets $(\text{PA}_0 - \text{PA}_{-1})$.

By substituting Expressions (5) through (8) back into Equations (3) and (4) and engaging in a bit of algebra, we can rewrite the difference between the informed and simple forecasts as:

$$\Delta \text{ Interest Cost} = (\Delta\text{PBO} * \text{DR}_{-1}) + (\text{PBO}_0 * \Delta\text{DR}) - \text{Realized Difference} \quad (9)$$

$$\Delta \text{ Expected Return on PA} = \Delta\text{PA} * \text{ERR}_0. \quad (10)$$

Each of the components in Equations (9) and (10) can be viewed as the forecasted change in interest cost and expected return on plan assets due to a specific change in pension plan parameters. The first term in Equation (9) is the anticipated change in following year's interest cost due to a change in the PBO from the previous year (Delta PBO). The second term is the anticipated change in interest cost due to a change in the discount rate (Delta Discount Rate), while the final term is the realized difference. In Equation (10) the entire expected change in next year's expected return is due to a change in the plan assets (Delta Plan Assets). This is because, unlike the discount rate, any change in the ERR will affect the current (year 0) earnings. The change in the *current* pension cost due to a change in the ERR can be estimated as:

$$\Delta \text{ Current Expected Return on PA (Delta ERR)} = PA_{-1} * \Delta ERR \qquad (11)$$

where ΔERR = Change in Expected Rate on Plan Assets, $(ERR_0 - ERR_{-1})$.

The decomposition of the interest cost and expected return on plan assets into the subcomponents detailed in Equations (9) through (11) are presented graphically in Figure 2.

Estimation of the Future Service Cost Component of Pension Cost

Estimating the expected change in the service cost is not as straightforward as forecasting the interest cost and expected return. The service cost is a function of the discount rate, the compensation rate, and various actuarial assumptions about a company's retirement obligations. To most accurately predict the effect of a change in the discount rate or compensation rate on future service cost, one would need to know the duration of the expected payments. Firms with longer durations should experience a larger change in service cost for a given change in either rate. Because companies are not required to report this value, I calculate a parameter that attempts to proxy for the relative duration of pension obligations. I then run an algorithm that estimates the change in each company's service cost based on the company's duration decile and the observed relation between rate changes and future service costs in years prior to the year of estimation. The details of this estimation are provided in Appendix A.

Calculation of Estimated Total Mechanical Change and Trading Portfolios

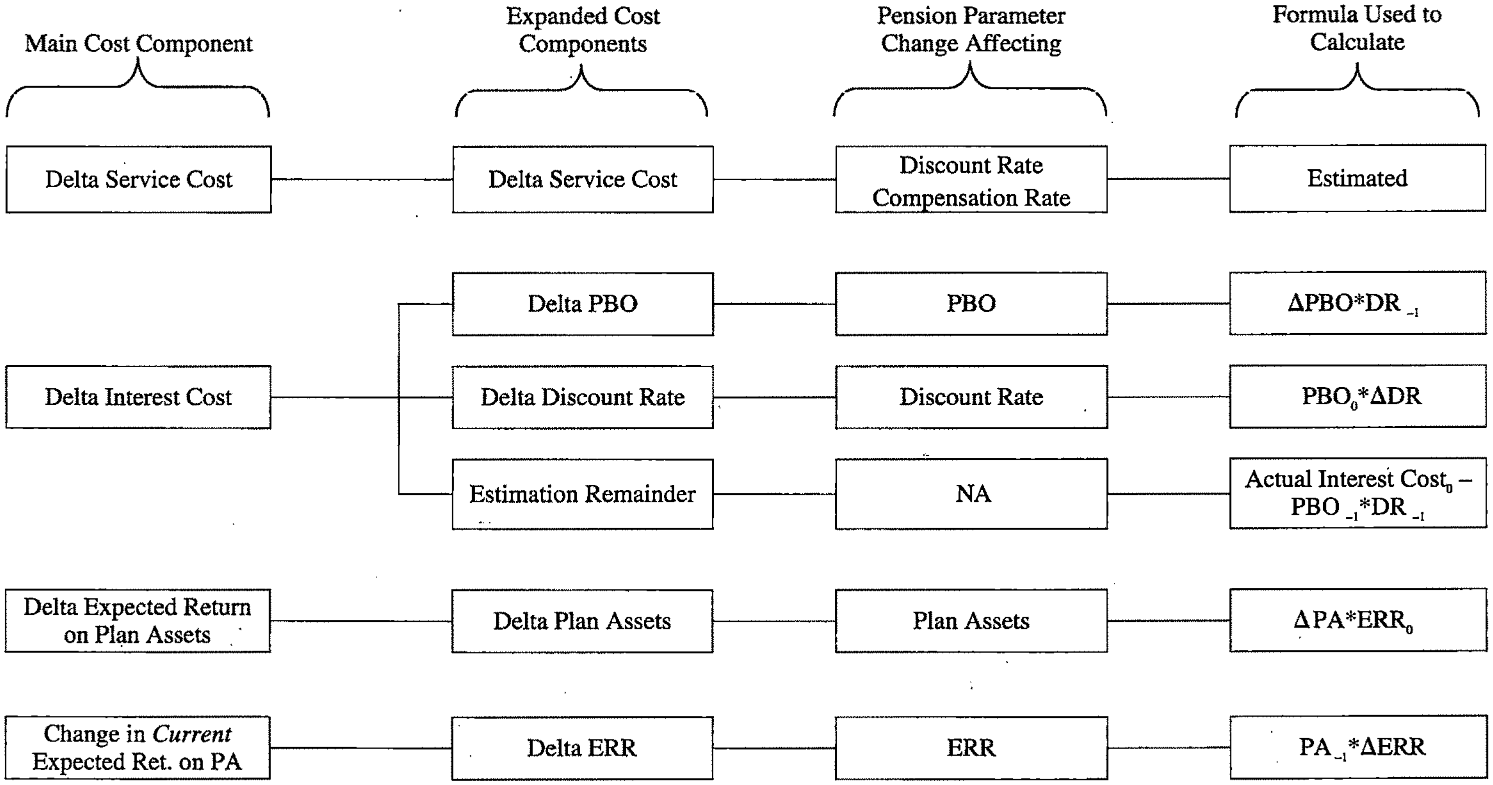
I estimate the total mechanical change in pension income (TOTMEC) as the sum of all the expanded pension cost components, excluding the interest cost remainder term and the Delta ERR term. Calculations of these components are detailed in Figure 2. These terms are summed in such a manner that they represent the total estimated increase in pension income (decrease in pension cost) in dollars per share due to changes in the discount rate, compensation rate, PBO, and plan assets. A negative value thus indicates that the firm is expected to have an increase in pension cost (i.e., a decrease in net income) due to pension plan parameter changes.

VI. RESULTS

Accuracy of "Informed" Pension Cost Component Forecasts

Table 2 examines whether the methodology I employ to make an informed estimate of future service cost, interest cost, and expected return on plan assets actually yields a more accurate estimate than simply assuming these costs will remain unchanged. Table 2 reports statistics for the absolute value of the difference between the actual service cost, interest

FIGURE 2
Breakdown of Estimated Changes in Pension Cost Components



where:
ΔDR = change in Discount Rate, $(DR_0 - DR_{-1})$;
ΔERR = change in Expected Rate of Return on Plan Assets, $(ERR_0 - ERR_{-1})$;
ΔPBO = change in PBO, $(PBO_0 - PBO_{-1})$; and
ΔPA = change in Plan Assets, $(PA_0 - PA_{-1})$.

TABLE 2
Differences in the Accuracy of the Simple and Informed Forecasts of Pension Cost Components

Absolute Differences between Actual and Forecasted Costs (simple and informed) (dollars per share)

1 Pension Cost Component	2 Rate Change	3 Forecast	4 n	5 Mean	6 Difference	7 t-statistic	8 Difference Upper 10%	9 Difference Upper 5%	10 Difference Top 1%
Service Cost	no	simple	1363	0.023	0.001	2.28**	>0.016	>0.024	>0.059
		informed	1363	0.022					
	yes	simple	3767	0.040	0.009	6.81***	>0.033	>0.053	>0.153
		informed	3767	0.031					
Interest Cost	no	simple	1974	0.032	0.008	5.86***	>0.034	>0.058	>0.158
		informed	1974	0.024					
	yes	simple	4700	0.056	0.019	10.38***	>0.052	>0.098	>0.328
		informed	4700	0.036					
Expected Return on Plan Assets	NA	simple	1940	0.098	0.013	2.98***	>0.070	>0.146	>0.582
		informed	1940	0.085					

, * Denotes significance at a probability <.05 and <.01, respectively, two-tailed tests.

Table 2 shows the accuracy of making an "informed" forecast of next year's (year 1's) pension cost components using the changes in the discount rate, compensation rate, PBO, and plan assets as opposed to a "simple" forecast that assumes all the pension cost components remain at their current level. The first column shows the cost component of interest. The second column shows whether a change in assumed rates has occurred (yes for service cost if there is a change in compensation rate or discount rate, yes for interest cost only if there is a change in discount rate). The fifth column gives the average absolute difference between the actual interest cost, service cost, and expected return on plan assets in the following year (year 1) and the simple and informed forecasts of these values based on year 0 pension costs and the information in the year 0 10-K. The sixth column highlights the difference between the average absolute error of the simple forecast and that of the informed forecast. The seventh column details whether the difference in the forecast accuracies is significant. The final columns show the lower bounds of the difference in the accuracy between the informed and simple models for firms where the informed model is a substantially better predictor than the simple model. All values are in dollars per share based on the shares outstanding at the 10-K release. Results for the service cost and interest cost are shown using data from 1993–2001. Results for the expected return on plan assets are shown using data from 1998–2001 because this information was not consistently available prior to the issuance of SFAS No. 132.

cost, and expected return on plan assets in year 1 and the simple and informed forecasts of these costs at year 0.[5] The second row shows that when there is a change in either the compensation rate or the discount rate, the informed forecast of service cost tends to be more accurate than the simple forecast by about .9 cents per share.[6] Error terms for informed forecasts also have a lower standard deviation than error terms for simple forecasts, but are not reported. If there are no rate changes, then the informed forecast of service cost is still slightly better than the simple forecast because it takes into account that employees are one year closer to retirement. When the informed method is used to predict next year's interest cost and there is no change in the discount rate, it performs, on average, about .8 cents better. When the discount rate does change, however, the informed forecast is, on average, 1.9 cents better than the simple forecast. Predicting the expected return on plan assets by adjusting for the change in plan assets also yields a significant improvement of 1.3 cents per share over the simple forecast. Additionally, when the informed forecast is substantially more accurate than the simple forecast, analysts could increase their forecast accuracy anywhere from 1.6 cents per share to well above 58 cents per share by incorporating the changes in the various pension parameters. These results demonstrate the effectiveness of the forecast methods employed.

Do Analysts Understand the Mechanical Impact of Current Period Pension Parameter Changes on Future Earnings?

Table 3 investigates whether analysts incorporate the new information contained in pension plan parameter changes into their forecasts of annual earnings. Following methodology employed by Rajgopal et al. (2003), Table 3 regresses monthly forecast errors on the forecasted pension cost changes and various year 0 financial variables.[7] The first column reports the results of regressing actual year 1 earnings on the financial variables and forecasted pension cost changes.[8] Column one shows that all forecasted cost changes (Delta Service Cost, Delta PBO, Delta Discount Rate, and Delta PA) have significant predictive ability for future earnings in the expected direction.[9] Additionally, the coefficients of the forecasted cost changes, with the exception of Delta Service Cost, are all close to 1. This means that, in general, a 1 cent forecasted change in pension cost results in a 1 cent change in future earnings. The −3.89 coefficient on Delta Service Cost indicates that future earnings tend to change even more in the anticipated direction than predicted by the forecast. This could be partially due to underestimation of the change in service cost or to other pension costs related to the service cost (in particular the amortization of pension gains or losses) that are not estimated in this paper due to Compustat limitations. Along with the

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[5] Table 2 uses data from 1993 to 2001 for the service cost and interest cost. For the expected return on plan assets Table 2 uses data from 1998 to 2001. This is because the actual expected return on plan assets is not available in Compustat until after 1998 when SFAS No. 132 went into effect.

[6] In results not shown, I find that approximately 70 percent of firms adjust their discount rate every year and that about 32 percent adjust their compensation rate. Additionally, about 20 percent of firms adjust their ERR in a given year, about half upward and half downward.

[7] Changes in the forecasted interest cost component are split into their causal elements as described in Section V, Equation (9).

[8] For these regressions, I include various control variables that have been shown to be predictive of future earnings or returns. This is based on past research showing a strong relationship between earnings and returns. The control variables included are the previous year's earnings, forecast error, book-to-market, size, earnings-to-price, beta, firm age, net operating assets, total accruals, the change in the S&P long bond yield, dummies indicating whether the firm missed or beat its forecast in the previous year, a dummy indicating whether the firm has a defined benefit pension plan, year dummies, and industry controls based on four-digit SIC codes. Due to space considerations, I do not report the coefficients for the control variables. They are available upon request.

[9] The coefficient of the interest cost realized difference, although marginally statistically significant out to month 3, has such a small magnitude as to make it economically irrelevant and is not discussed in any further detail.

TABLE 3
Regressions of Monthly Analyst Forecast Errors on the Estimated Earnings Effects of Pension Plan Changes

			Forecast Month Relative to the Estimated Release of the Financial Statements for the Year Prior to Year of Predicted Earnings									
Regression Parameter	**Exp. Sign**	**On Future Earnings**	**Month −1**	**Month 0**	**Month 1**	**Month 2**	**Month 3**	**Month 4**	**Month 5**	**Month 6**	**Month 7**	**Month 8**
Intercept	?	0.481	0.165	0.189	0.129	0.079	0.072	0.066	0.056	0.045	0.013	0.035
		6.50***	−1.84*	−2.06**	−1.38	−0.93	−0.88	−0.88	−0.77	−0.64	−0.20	−0.53
Delta Service Cost	−	−3.889	−4.507	−4.303	−4.538	−3.916	−2.751	−1.379	−1.530	−1.149	−0.875	−0.587
		−9.63***	−12.16***	−11.49***	−10.91***	−10.01***	−7.35***	−3.94***	−4.50***	−3.45***	−2.72***	−1.86**
Delta PBO	−	−1.361	−2.962	−2.607	−2.089	−2.020	−1.792	−1.372	−1.223	−0.913	−0.469	−0.267
		−9.95***	−18.25***	−16.38***	−11.85***	−12.26***	−11.35***	−9.31***	−8.56***	−6.52***	−3.47***	−2.01**
Delta Disc. Rate	−	−0.904	−1.888	−1.683	−1.608	−1.458	−1.238	−1.057	−1.050	−0.742	−0.504	−0.348
		−4.30***	−9.47***	−8.58***	−7.22***	−7.04***	−6.26***	−5.70***	−5.85***	−4.22***	−2.96***	−2.09**
Interest Cost Realized Diff.	?	0.001	0.001	0.001	0.001	0.001	0.001	0.000	0.000	0.000	0.000	0.000
		1.84*	2.15**	1.82*	1.57	1.82*	1.85*	1.28	1.07	1.00	0.23	0.20
Delta Plan Assets	+	1.322	2.055	1.719	1.374	1.376	1.206	0.922	0.831	0.646	0.361	0.219
		13.89***	19.60***	15.89***	12.15***	12.87***	11.79***	9.67***	8.97***	7.12***	4.12***	2.55***
Delta ERR	?	−1.869	−0.930	−0.642	−0.437	−0.496	−0.564	−0.543	−0.635	−0.630	−0.664	−0.558
		−7.83***	−3.54***	−2.51**	−1.47	−1.77*	−2.11**	−2.17**	−2.61***	−2.64***	−2.88***	−2.46**
Actual Earnings Previous Year	?	0.798	−0.196	−0.184	−0.173	−0.164	−0.139	−0.101	−0.080	−0.062	−0.044	−0.038
		104.31***	−23.50***	−21.29***	−19.12***	−19.41***	−17.23***	−13.44***	−10.98***	−8.69***	−6.37***	−5.62***
Number of observations		21,459	11,534	10,307	14,146	15,366	15,769	16,062	16,276	16,369	16,465	16,544
R^2 w/o Pension Variables		0.550	0.100	0.101	0.057	0.054	0.044	0.030	0.025	0.021	0.016	0.013
Full Model R^2		0.559	0.152	0.149	0.081	0.072	0.054	0.032	0.027	0.021	0.014	0.012

(*continued on next page*)

TABLE 3 (Continued)

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*, **, *** Denotes significance at a probability of <.10, <.05, and <.01, respectively.

Tests for all parameters with expected signs are one-tailed tests. All t-tests with no expected sign are two-tailed tests.

Table 3 shows the results of pooled regressions of monthly analyst forecast errors on the informed forecasts of pension cost components and various control variables. The first column shows a regression of actual future earnings on the informed forecasts and the control variables. The form of the regression for months −1 to 8 is:

$$\begin{aligned}\text{Consensus Analyst Forecast Error}_{it} = {} & \alpha + \beta_1 * (\text{Delta Service Cost}) + \beta_2 * (\text{Delta PBO}) + \beta_3 * (\text{Delta Discount Rate}) + \beta_4 * (\text{Interest Cost Realized Diff}) \\ & + \beta_5 * (\text{Delta Plan Assets}) + \beta_6 * (\text{Delta ERR}) + \beta_7 * (\text{Previous Year's Earnings}) \\ & + \beta_8 * (\text{Previous Year's Forecast Error}) + \beta_9 * (\Delta\text{Long Bond Yield}) + \beta_{10} * \text{Miss} + \beta_{11} * \text{Beat} + \beta_{12} * \text{B/M} \\ & + \beta_{13} * \text{Size} + \beta_{14} * \text{Beta} + \beta_{15} * \text{E/P} + \beta_{16} * \text{Age} + \beta_{17} * \text{Net Operating Assets} + \beta_{18} * \text{Total Accruals} \\ & + \beta_{19} * \text{Pension Plan Dummy} + \text{Industry Dummies} + \text{Year Dummies}\end{aligned}$$

Forecast error is defined as actual future earnings minus the constructed consensus forecast for the given month. To preclude stale forecasts from biasing the consensus forecast the consensus forecast in month −1 and month 0 is generated using only forecasts made in those particular months. For months 1 to 8, I generate the consensus forecast based on the most recent forecast available from each analyst provided the forecast was made in month 1 or later. The change in the interest cost is broken into three components in the regression: a change due to a change in the plan benefit obligation (Delta PBO), a change due to a change in discount rate (Delta Discount Rate), and a realized difference not attributable to either (Interest Cost Realized Diff). The estimated change in the future expected return on plan assets is entirely due to changes in a firm's plan assets (Delta ERR). Delta ERR is the change in current earnings due to a new ERR assumption. Miss and beat are dummy variables taking a value of 1 if the firm missed or beat the earnings forecast in the previous year. Previous year's forecast error is the amount (in dollars per share) the firm missed or beat its annual consensus forecast by in the previous year. ΔLong Bond Yield is the change in the treasury long bond yield in the previous month and is included since it is highly correlated with changes in pension discount rates. Firm age is included as a control variable due to its relationship to the duration of the outstanding pension liability. Size, book-to-market, Beta, earnings-to-price, total accruals scaled by lagged total assets and net operating assets scaled by lagged total assets are also included in the regression as other variables since these variables have been shown in prior literature to be predictive of future returns and, due to the strong relationship between earnings and returns, may also prove predictive of future earnings.

See Appendix B for a more detailed description of the control variables.

All control variables, except total accruals and net operating assets are included as their respective decile rank scaled to range from a value of 0 to 1. Ranks are assigned to the variables of all December year-end firms and the variables for firms with January to November year-ends in the following year are ranked based on the cutoffs from the December ranking. This ensures that all rankings are done using data available at the date of ranking. Ranking based on fiscal year yields no qualitative difference in the results. Total accruals and net operating assets are included as their respective quintile rank scaled from 0 to 1. Age is calculated as the difference between a firm's fiscal year and the fiscal year of its earliest valid CRSP data. A pension plan dummy variable is included and takes on the value of 1 if the firm has a defined benefit pension plan. Eleven industry dummies based on firms' four-digit SIC codes are also included in the regression to control for industry effects. Year dummies are also included to control for possible yearly effects.

The pooled regressions use data from 1993–2001 since rate data is not available on Compustat until 1991 and two years of data are required for service cost forecasts.

results in Table 2, this regression shows that changes in the noted pension parameters are reliable predictors of changes in future earnings and, thus, worth the attention of both analysts and investors.[10]

The remaining columns of Table 3 explore H1 by regressing the analyst consensus forecast error in the months relative to the 10-K release on the forecasted changes in pension cost. Because the exact date of the 10-K release is unknown, I assume that it is released sometime during the third month after the fiscal year-end. This month is designated month 0. Although most analysts probably have access to the 10-K information by the time of the month 0 consensus forecast, I conservatively assume that all analysts do not have the information until month 1. I generate the consensus analyst forecast for each month using individual analyst forecast data available from I/B/E/S. To preclude stale forecasts from biasing the consensus forecast I do the following: The consensus forecast in month −1 and month 0 is generated using only forecasts made in those particular months. This ensures that I only pick up the effects of information available in those particular months. For months 1 to 8, I generate the consensus forecast based on the most recent forecast available from each analyst, provided the forecast was made in month 1 or later. This assures that all analysts included in the consensus forecast have access to the 10-K pension information when making their forecasts.

The month −1 column checks to see if analysts have obtained private information prior to the assumed 10-K release date and, thus, already incorporated the pension change information into their forecasts. The results can be interpreted as the forecast error, in cents, associated with a 1 cent estimated increase in pension cost due to changes in the specified pension parameters. For instance, in month −1, an estimated 1 cent increase in service cost results in analysts overestimating earnings by 4.5 cents. As this column shows, the four pension cost change estimates all significantly predict analyst forecast error in the expected direction and with coefficients of magnitude similar to or greater than those in column one. Analysts therefore do not seem to have incorporated the pension information into their earnings forecasts prior to the release of the financial statements.

In months 0 and 1 there is no significant change in the magnitude or significance of the pension change coefficients from month −1, even though the information is publicly available to most analysts. Besides the 10-K information, firm's first quarter earnings announcements should also be available during month 1. These earnings should reflect the new pension information and, thus, serve as another aid for analysts to update their forecasts. Consistent with H1, however, analysts do not appear to incorporate the pension change information into their forecasts at the time it becomes publicly available.

The remaining columns show the effect of pension changes on the forecast errors in subsequent months. Although the magnitude of the coefficients steadily declines over time, the table indicates that analysts do not seem to appreciably process the earnings implications of pension plan changes until at least the fourth month after the release of the 10-K. This is about the time that the second quarter's earnings report is issued and incorporated into analysts' earnings forecasts. Even with the partial incorporation of pension information in the fourth month, the pension cost change components remain significant predictors of analyst forecast error into the eighth month. These results are consistent with H1 and imply that analysts do not specifically incorporate the information contained within the pension notes, but rather realize the effects of pension cost changes gradually through the release of subsequent quarterly earnings announcements.

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[10] Column 1 also shows that delta ERR has a statistically significant coefficient of −1.87, even though changes in ERR should mechanically have no predictive power for future earnings. This is consistent with management raising the ERR in an effort to boost current earnings and mask poor performance, although this hypothesis is not specifically tested in the paper.

Table 4 reinforces the conclusions drawn from Table 3 by examining whether the combined changes in pension parameters are predictive of future analyst forecast errors of quarterly earnings at the time of announcement. Table 4 sums all of the individual pension cost change components in Table 3 (excluding Delta ERR) to arrive at a single variable, TOTMEC, which indicates the total estimated change in future pension cost based on the 10-K information. Regressing analyst forecast errors of quarterly earnings on TOTMEC reveals that this variable is predictive of forecast errors in the three quarters following the release of the pension change information. This is consistent with H2 and once again demonstrates that analysts do not fully incorporate the pension change signals into their forecasts. It is interesting to note that in the fourth quarter TOTMEC is predictive of forecast error in the opposite direction. This indicates that analysts may, in fact, overreact to their errors in the previous three quarters when forecasting annual earnings during the fourth quarter.

The evidence in Tables 3 and 4 is consistent with H1 and H2 and demonstrates that analysts do not fully incorporate relevant and publicly available information contained in the pension notes at the time of its release, even though they deal with this information on a repeated basis. This may be due to the complexity of the information presented and the effort costs associated with incorporating this information. More detailed disclosure on the effects of rate changes on pension costs and PBO might make this information more accessible both to analysts and investors. In particular, sensitivity disclosures similar to those required for OPEBs (including the effect of rate changes on service cost) and the disclosure of an aggregated duration for pension liabilities would significantly reduce the effort required to incorporate the earnings implications of pension plan parameter changes. The requirement in the SFAS No. 132 revision to report an estimated payment schedule for projected benefits will facilitate duration calculations and should, therefore, improve the accuracy of service cost estimates.

Do Investors Understand the Mechanical Impact of Current Period Pension Parameter Changes on Future Earnings?

While it is clear from Tables 3 and 4 that analysts fail to readily incorporate pension information into their forecasts, it remains to be seen whether the market likewise fails to expeditiously process this information. To investigate this question, I examine whether a simple trading strategy based on TOTMEC is able to generate future abnormal returns. Firms are first sorted into TOTMEC quintile ranges by calendar year. All firms in the middle three quintiles and any firms that do not have a December fiscal year-end are eliminated. The December fiscal year-end requirement ensures that the trading strategy can be easily implemented. Firms in the upper TOTMEC quintile that do not have an estimated gain in income of at least 1 cent and those in the bottom quintile that do not have an estimated loss of at least 1 cent are also eliminated. This ensures the strength of the trading strategy signals. I then take a long position in the remaining upper quintile firms and a short position in the bottom quintile firms beginning April 1st of the following year. This ensures that the 10-K information is available. Portfolio formation begins in 1994 because 1993 is the first year that I can compute all pension cost change components. It ends in 2002 due to the availability of returns data.

I check to see whether the TOTMEC signal is in fact a good predictor of future pension costs. The high TOTMEC portfolio has an average decrease in pension costs of 8 cents, while the low TOTMEC portfolio has an average increase in pension costs of 11 cents in

TABLE 4
Regressions of Quarterly Earnings Forecast Error on Total Estimated Mechanical Change in Pension Income

Regression Parameter	Quarter 1	Quarter 2	Quarter 3	Quarter4
Intercept	−0.012	0.008	0.005	0.005
	−0.92	0.74	0.44	0.30
TOTMEC	0.124	0.175	0.045	−0.095
	8.42***	14.32***	3.39***	−4.65
Missed Forecast in Previous Quarter	−0.012	−0.029	−0.036	−0.045
	−2.49**	−8.60***	−10.22***	−8.28***
Beat Forecast in Previous Quarter	0.013	0.023	0.031	0.016
	2.92***	7.05***	9.20***	3.01***
Book-to-Market	−0.010	−0.003	−0.005	−0.009
	−1.77*	−0.78	−0.99	−1.15
Size	−0.028	−0.006	−0.021	−0.052
	−4.15***	−1.10	−3.53***	−5.75***
Momentum	0.038	0.018	0.015	0.015
	9.15***	5.31***	4.10***	2.62***
Earnings-to-Price	0.029	0.006	0.006	0.050
	5.53***	1.35	1.40	7.10***
Beta	−0.001	−0.003	−0.010	−0.011
	−0.10	−0.81	−2.12**	−1.55
Age	−0.019	−0.006	−0.013	0.002
	−3.74***	−1.49	−2.89***	0.28
Net Operating Assets	−0.008	−0.011	−0.018	−0.039
	−1.91*	−3.31***	−5.00***	−7.14***
Total Accruals	−0.012	−0.012	−0.007	−0.019
	−3.07***	−3.78***	−1.96***	−3.47***
Industry Dummies	Included	Included	Included	Included
Year Dummies	Included	Included	Included	Included

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*, **, *** Denotes significance at a probability of <.10, <.05, and <.01, respectively.
The t-tests for TOTMEC are one-tailed tests. All other tests are two-tailed tests.
Table 4 shows the results of pooled regressions of future quarterly earnings forecast errors on total estimated mechanical change in pension income. The regressions use data from 1993–2000. The regressions take the form:

$$\begin{aligned}\text{Quarterly Earnings Forecast Error}_t = {} & \alpha + \beta_1 * \text{TOTMEC} + \beta_2 * \text{Miss}_{t-1} + \beta_3 * \text{Beat}_{t-1} + \beta_4 * \text{B/M} \\ & + \beta_5 * \text{Size} + \beta_6 * \text{Momentum} + \beta_7 * \text{E/P} + \beta_8 * \text{Beta} \\ & + \beta_9 * \text{Age} + \beta_{10} * \text{Net Operating Assets} \\ & + \beta_{11} * \text{Total Accruals} + \text{Industry Dummies} + \text{Year Dummies}.\end{aligned}$$

TOTMEC, the variable of interest, is the sum of all the expanded pension cost components, excluding the Interest Cost Realized Diff. and Delta ERR terms, which are detailed in Figure 2 and described in Section V. These terms are summed in such a manner that they represent the total estimated increase in pension income (decrease in pension cost) in dollars per share due to changes in the PBO, plan assets, discount rate, and compensation rate. All other variables in the regression are included as controls.
Quarterly earnings forecast error is calculated as the firm's actual quarterly earnings minus the analyst consensus forecast at the earnings announcement date. Momentum is the six-month, market-adjusted return ending one month prior to the end of the firm's fiscal year. It is included as its respective decile rank scaled between 0 and 1. It is used as a control because prior literature has shown that it is predictive of future returns and thus may also prove predictive of future earnings surprises.
All other control variables are as described in Table 3.

the following year. The long and short portfolios thus have an average difference of 19 cents in pension cost change. It is this predictable cost difference that drives the Table 5 results.

Panel A of Table 5 shows the yearly returns for each portfolio by their year of formation and the return to the long-short hedge portfolio. The hedge portfolio generates positive returns in seven of the nine years and in the two "off" years it loses less than 1 percent. Over the nine years, average portfolio return is 4.6 percent. The magnitude and consistency of this return demonstrates that the pension change information is worth consideration by both investors and analysts. In order to ensure that the results are not driven by some other factor correlated with TOTMEC, I estimate cross-sectional regressions of yearly size-adjusted returns on a TOTMEC portfolio dummy and various other control variables known to be predictive of future returns. Panel B of Table 5 shows the average coefficients of these regressions. The results indicate a statistically significant average hedge return of 7.8 percent, indicating the robustness of the TOTMEC signal. Panels A and B are consistent with H3 and demonstrate that, like analysts, investors also fail to expeditiously incorporate changes in disclosed pension information. Instead, both these groups seem to gradually incorporate this information by observing its effects on quarterly earnings. Because all this information is publicly available, information assimilation would appear to drive this inefficiency.

Fama-MacBeth Regressions of Future Returns on Pension Parameters

Table 6 examines how the market responds to the magnitude of six pension parameters that may influence future long-run earnings and cash flows. Both panels of Table 6 show the results of multivariate Fama-MacBeth regressions of the cumulative size-adjusted one- to five-year future returns beginning the month after the estimated 10-K release on selected pension parameters and various control variables shown in previous literature to be predictive of future returns. All variables are ranked into either deciles or quintiles and their ranks then scaled between 0 and 1. This allows the regression coefficients to be interpreted as the returns associated with taking a long position in the highest decile or quintile of each variable and taking a short position in the lowest decile or quintile.[11]

Panel A of Table 6 tests H4, H5a, and H5b by regressing future returns on the control variables and three pension liability measures: the PBO and the on- and off-balance-sheet liabilities, all scaled by total assets. The results show that the PBO is predictive of future returns going out five years and that the off-balance-sheet liability is predictive of future returns in years 2 through 5, both in the expected negative direction. These findings are consistent with H4 and H5b. In both cases the magnitude of the cumulative returns associated with the pension parameters increases steadily over the five-year window, with PBO yielding a return differential of approximately 25 percent and the off-balance-sheet liability yielding a return differential of 15 percent over the five-year period. In keeping with H5a, however, on-balance-sheet liabilities are not predictive of cumulative future returns in any of the following five years. This is consistent with investors readily incorporating information that has already rolled through the income statement and is currently

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[11] The pension liability parameters (PBO and on- and off-balance-sheet liabilities) are ranked in deciles due to the wide dispersion in these values. The pension rate parameters are ranked in quintiles because the rates are clustered in a limited range and this avoids empty ranks. All the control variables are ranked in deciles except net operating assets (NOA) and total accruals, which are ranked in quintiles. This is in keeping with prior literature. Ranking NOA and total accruals in deciles does not qualitatively affect the results.

TABLE 5
Returns Associated with Total Estimated Mechanical Change Portfolios

Panel A: Yearly Results of Hedge Portfolios Formed Based on TOTMEC

	Year of Portfolio Formation									
	1994	**1995**	**1996**	**1997**	**1998**	**1999**	**2000**	**2001**	**2002**	**Average**
Income-Increasing Pension Changes	0.154	0.294	0.126	0.443	−0.187	0.116	0.100	0.225	−0.286	0.109
Income-Decreasing Pension Changes	0.025	0.256	0.123	0.370	−0.183	0.071	−0.021	0.206	−0.279	0.063
Yearly Long-Short Portfolio Returns	0.129	0.038	0.003	0.073	−0.004	0.045	0.121	0.019	−0.007	0.046
Firms in Long Portfolio	112	42	132	111	118	99	124	54	5	88.6
Firms in Short Portfolio	92	125	86	58	53	112	53	118	108	89.4
Total Number of Firms	204	167	218	169	171	211	177	172	113	178.0

Panel B: Yearly Cross-Sectional Regressions of One-Year, Size-Adjusted Returns on TOTMEC

The form of the regression is:

$$R_{it} = \alpha + \beta_1 * \text{TOTMEC Portfolio} + \beta_2 * \text{Previous Year's Forecast Error} + \beta_3 * \text{B/M} + \beta_4 * \text{Momentum} + \beta_5 * \text{E/P} + \beta_6 * \text{Beta} + \beta_7 * \text{Age} + \beta_8 * \text{Net Operating Assets} + \beta_9 * \text{Total Accruals} + \text{Industry Dummies}$$

Parameter	Expected Sign	Annual Average Coefficient	t-statistic
Intercept	?	0.009	0.10
TOTMEC Portfolio	+	0.078	3.46***
Previous Year's Forecast Error	+	0.044	1.45*
Book-to-Market	+	0.021	1.66*
Momentum	?	0.008	0.13
Earnings-to-Price	+	−0.009	−0.09
Beta	?	−0.073	−2.11
Age	?	0.010	0.15
Net Operating Assets	−	−0.021	−0.36
Total Accruals	−	−0.176	−3.71***

(*continued on next page*)

TABLE 5 (Continued)

*, *** Denotes significance of a probability of <.10 and <.01, respectively, in two-tailed tests.

TOTMEC is the sum of all the expanded pension cost components, excluding the Interest Cost Realized Diff. and Delta ERR terms, which are detailed in Figure 2 and described in Section V. These terms are summed in such a manner that they represent the total estimated increase in pension income (decrease in pension cost) in dollars per share due to changes in the PBO, plan assets, discount rate, and compensation rate.

Estimated pension income-increasing TOTMEC portfolios are generated by dividing firms into TOTMEC quintiles by calendar year. Firms in the middle three quintiles and any firms that do not have a December fiscal year-end are then eliminated. The December fiscal year-end requirement ensures that the trading strategy can be easily implemented. Firms in the upper TOTMEC quintile that do not have an estimated gain in income of at least 1 cent and those in the bottom quintile that do not have an estimated loss of at least 1 cent are also eliminated. This ensures the strength of the trading strategy signals. A long position is taken in the remaining upper quintile firms and a short position in the bottom quintile firms beginning April 1st of the following year. This ensures that the 10-K information is available. Portfolio formation begins in 1994 since 1993 is the first year that all pension cost change components can be calculated. Formation ends in 2002 due to the availability of returns data. The results of holding the long and short portfolios a year from formation are reported.

Panel B shows the average coefficients of yearly cross-sectional regressions of future one-year, size-adjusted returns on the total estimated mechanical change in earnings due to pension parameter changes (TOTMEC). The TOTMEC Portfolio variable takes on a value of 1 if a firm is in the high TOTMEC portfolio and a value of 0 if it is in the low TOTMEC portfolio.

All control variables are as described in Tables 3 and 4.

TABLE 6
Regressions of Future Returns on Various Pension Parameters

The yearly regressions in Panel A take the form:

$$\begin{aligned}R_{it} = {} & \alpha + \beta_1 * \text{PBO} + \beta_2 * \text{On-Balance-Sheet Liability} \\ & + \beta_3 * \text{Off-Balance-Sheet Liability} + \beta_4 * \text{Previous Year's Forecast Error} \\ & + \beta_5 * \text{B/M} + \beta_6 * \text{Size} + \beta_7 * \text{E/P} + \beta_8 * \text{Beta} + \beta_9 * \text{Age} \\ & + \beta_{10} * \text{Net Operating Assets} + \beta_{11} * \text{Total Accruals} + \text{Industry Dummies}\end{aligned}$$

		Cumulative Size-Adjusted Returns				
Parameter	**Expected Sign**	**1-Year**	**2-Year**	**3-Year**	**4-Year**	**5-Year**
Panel A: Regression of Future Returns on Pension Liability Measures						
Intercept	?	0.056 (0.95)	0.187 (2.12*)	0.274 (3.54***)	0.244 (1.96*)	0.393 (2.62**)
Plan Benefit Obligation	−	−0.067 (−3.49***)	−0.133 (−4.1***)	−0.191 (−2.75***)	−0.203 (−2.99***)	−0.246 (−3.26***)
On-Balance-Sheet Liabilities	?	−0.008 (−0.82)	−0.004 (−0.21)	−0.011 (−0.53)	−0.018 (−0.82)	−0.034 (−1.46)
Off-Balance-Sheet Liabilities	−	−0.018 (−1.20)	−0.048 (−2.28**)	−0.078 (−3.47***)	−0.106 (−3.97***)	−0.147 (−3.46***)
Previous Year's Forecast Error	+	0.027 (1.85*)	0.056 (3.17***)	0.082 (3.54***)	0.081 (2.87***)	0.075 (1.64*)
Book-to-Market	+	0.03 (1.46*)	0.045 (0.75)	0.121 (1.20)	0.218 (1.64*)	0.395 (2.81**)
Momentum	?	0.07 (2.18**)	0.024 (0.63)	0.007 (0.14)	0.033 (0.51)	0.07 (0.53)
Earnings-to-Price	+	0.067 (1.61*)	0.078 (0.90)	0.041 (0.44)	0.028 (0.53)	0.009 (0.10)
Beta	?	−0.027 (−1.25)	−0.057 (−1.46)	−0.082 (−1.97*)	−0.088 (−1.30)	−0.066 (−0.68)
Age	?	0.009 (0.33)	0.002 (0.04)	0.016 (0.19)	0.059 (0.62)	−0.025 (−0.20)
Net Operating Assets	−	−0.076 (−5.12***)	−0.095 (−3.58***)	−0.114 (−3.09***)	−0.071 (−1.41*)	−0.127 (−2.65**)
Total Accruals	−	−0.12 (−8.05***)	−0.179 (−6.24***)	−0.217 (−6.28***)	−0.284 (−6.41***)	−0.397 (−4.52***)

Panel B: Regression of Future Returns on Pension Plan Rate Assumptions

The yearly regressions take the form:

$$\begin{aligned}R_{it} = {} & \alpha + \beta_1 * \text{ERR} + \beta_2 * \text{Discount Rate} + \beta_3 * \text{Compensation Rate} \\ & + \beta_4 * \text{Previous Year's Forecast Error} + \beta_5 * \text{B/M} + \beta_6 * \text{Size} + \beta_7 * \text{E/P} \\ & + \beta_8 * \text{Beta} + \beta_9 * \text{Age} + \beta_{10} * \text{Net Operating Assets} + \beta_{11} * \text{Total Accruals} \\ & + \text{Industry Dummies}\end{aligned}$$

Intercept	?	−0.035 (−0.63)	0.046 (0.85)	−0.003 (−0.04)	−0.163 (−0.71)	−0.236 (−0.93)
Expected Rate of Return on Plan Assets	?	0.037 (1.68)	0.03 (0.71)	0.046 (1.78)	0.054 (1.36)	−0.01 (−0.19)

(*continued on next page*)

TABLE 6 (Continued)

Parameter	Expected Sign	Cumulative Size-Adjusted Returns				
		1-Year	2-Year	3-Year	4-Year	5-Year
Panel B: Regression of Future Returns on Pension Plan Rate Assumptions (Continued)						
Discount Rate	?	−0.009	−0.043	0.019	0.067	0.132
		(−0.28)	(−0.76)	(0.25)	(0.97)	(1.66)
Compensation Rate	?	0.028	0.032	0.068	0.12	0.276
		(2.16*)	(1.37)	(1.96*)	(4.07***)	(6.49***)
Previous Year's Forecast Error	+	0.034	0.044	0.061	0.05	−0.033
		(1.84**)	(2.1**)	(1.94**)	(1.03)	(−0.37)
Book-to-Market	+	0.055	0.064	0.134	0.234	0.529
		(2.79***)	(0.89)	(1.03)	(1.37)	(2.94**)
Momentum	?	0.053	0.012	−0.018	0.079	0.227
		(1.24)	(0.20)	(−0.17)	(0.61)	(0.85)
Earnings-to-Price	+	0.068	0.093	0.049	−0.026	−0.066
		(1.12)	(0.87)	(0.41)	(−0.37)	(−0.89)
Beta	?	−0.038	−0.087	−0.115	−0.119	−0.078
		(−1.49)	(−1.66)	(−3.61***)	(−1.12)	(−0.56)
Age	?	0.007	0.018	0.08	0.136	0.019
		(0.25)	(0.40)	(1.26)	(1.30)	(0.16)
Net Operating Assets	−	−0.045	−0.047	−0.056	0.037	0.016
		(−2.43**)	(−1.17)	(−1.15)	(0.33)	(0.15)
Total Accruals	−	−0.128	−0.185	−0.247	−0.319	−0.481
		(−6.32***)	(−5.59***)	(−5.29***)	(−5.21***)	(−5.04***)

*, **, *** Denotes significance at a probability of <.10, <.05, and <.01, respectively. All tests with an expected sign are one-tailed test. All tests with an unknown expected sign are two-tailed tests.

Due to the relatively small number of years with relevant pension data, the Newey-West correction may not be accurate beyond the two-year return window. For years 3 to 5, therefore, the lower of the Newey-West or the unadjusted t-statistic is reported.

Table 6 shows the results of Fama-MacBeth cross-sectional type regressions of future size-adjusted cumulative returns on various pension parameters. Results are shown for returns in the one- to five-year time periods beginning in the month following the estimated month in which the firm's 10-K is released. Yearly regressions are performed by calculating the pension parameters and control variables in a given fiscal year and then regressing future returns on these variables. The coefficients of these yearly regressions are then averaged to produce a single coefficient and statistical significance is assessed using Newey-West corrected t-statistics based on the yearly coefficients. The columns show the average coefficients (interpreted as cumulative portfolio returns) for the given horizon. t-statistics are show below in parentheses.

Panel A evaluates the effect of the firm's total pension liability (PBO) and the on- and off-balance-sheet components of the firm's unfunded pension liability on future returns. This panel uses data from 1988–2001.

Panel B tests whether the various pension plan rate assumptions are predictive of future returns. This panel uses data from 1991–2001 due to Compustat availability.

R_{it} is the cumulative return in the one- to five-year window beginning in the month after the estimated month of release of firm i's annual report. The pension parameters in Panel A are sorted into decile ranks and the pension parameters in Panel B are sorted into quintile ranks. Quintiles are used in Panel B due to the relatively tight clustering of the pension rates. Using quintiles helps avoid empty ranks. Ranks are assigned to the variables of all December year-end firms and the variables for firms with January to November year-ends in the following year are ranked based on the cutoffs from the December ranking. This ensures that all rankings are done using data available at the date of ranking. All ranks are scaled from 0 to 1 to allow the regression coefficients to be interpreted as the return differential between the high and low ranks.

The pension parameters used are defined in Appendix B.

All control variables are as defined in Tables 3 and 4.

reflected on the balance sheet, but failing to fully account for the less salient off-balance-sheet liability buried in the pension notes. Investors likewise do not appear to fully appreciate the future cash flow and earnings implications associated with a large total pension liability, the majority of which is also not reflected on the balance sheet.[12]

Panel B of Table 6 tests H6 by regressing future returns on the control variables and levels of the three disclosed pension rates: the expected rate of return on plan assets, the discount rate, and the compensation rate. Neither the ERR nor the discount rate assumptions are predictive of future returns. I cannot therefore reject the null hypotheses of H6 for either of these rates. This implies that the market is either able to adjust for the earnings effects of high rate assumptions or that the earnings benefits of these aggressive rate assumptions are cancelled out by correlated negative earnings information. It may also be that firms are unable or unwilling to maintain abnormally high rate assumptions for extended periods.

Panel B of Table 6 also shows that firms adopting low compensation rates tend to experience significantly lower cumulative returns over the next five years than firms reporting high compensation rates (although the cumulative return is not statistically significant in the second year). Indeed, over a five-year period, firms with the highest compensation rates experience a 28 percent return differential over their counterparts with low compensation rates. This could indicate that the compensation rate is a relatively unbiased bellwether for managerial optimism because the negative earnings consequences of choosing a high rate would preclude managers from being overly optimistic in this assumption. A high compensation rate may therefore signal that management expects strong future performance and attendant generous wage increases. The null hypothesis for H6 can, therefore, be rejected for the compensation rate.

VII. CONCLUSION

This paper explores whether investors and analysts fully process publicly available pension information when establishing prices and making earning forecasts. I examine whether investors and analysts expeditiously process the readily estimable earnings effects of changes in the PBO, plan assets, discount rate, and compensation rate. I find that both prices and forecasts fail to reflect new pension information at the time it becomes publicly available and that it is only gradually incorporated through its effects on quarterly earnings. These findings provide insight into how analysts form forecasts and to what level they scrutinize and incorporate disclosed information. In particular, I demonstrate that analysts fail to incorporate the quantifiable earnings effects of relevant and economically significant information even when they deal with this information on a repeated basis and it can be reliably estimated on a per-firm basis.

This paper also investigates whether investors have trouble assessing the long-run cash flow and earnings implications associated with the magnitude of the PBO, the on- and off-balance-sheet liabilities and the three major pension rate assumptions. I find that both the PBO and the off-balance-sheet liability are predictive of future returns over the next five years, but the on-balance-sheet liability is not. This suggests that investors can completely process the pension information that has already been recognized in income, but fail to fully impound the valuation impact of pension liabilities disclosed only in footnotes. I find that higher compensation rates are predictive of higher future returns, but the ERR and the discount rate are not. Management's choice of compensation rate may therefore signal their expectation for future firm performance.

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[12] In separate, unreported regressions I find that the one-year returns for changes in the off-balance-sheet liability and PBO are also significant and in the expected direction.

By identifying several economically significant parameter changes that are not fully reflected in prices and forecasts, this paper offers a benchmark to judge whether the SFAS No. 132 revision has improved the quality of pension information provided to investors and analysts. These findings should also be helpful in highlighting issues that the IASB and FASB might consider in their deliberations on setting mutually acceptable pension accounting standards. The explanation of why analysts and investors fail to incorporate specific pension information into prices and forecasts remains an open question. To gain a better understanding of the process, future research needs to deal with the issues of current pension accounting complexity and the fact that most of the economically significant pension information is buried in lengthy footnotes.

APPENDIX A

Service Cost Estimation

For firms reporting ABO, I use the ABO divided by the PBO as a proxy for pension plan duration. The rationale behind this is that the larger a firm's ABO relative to its PBO, the closer its workforce is to retirement and, thus, the lower the service cost's duration. For firms that do not report ABO, I use the ratio of interest cost divided by service cost plus interest cost as an alternate proxy. The rationale behind this is similar to that used for the ABO to PBO ratio.[13] I then sort firms into duration deciles by year based on the appropriate proxy.[14]

I assume that a 1 percent change in the discount (compensation) rate will cause an n percent (m percent) change in the service cost of the "average duration" firm, with both n and m increasing (decreasing) for firms in the lower (higher) duration deciles.[15] My methodology essentially entails simultaneously estimating the values of the n, m, and associated shift estimators (based on firm-specific duration deciles) that produce the smallest combined absolute error between my estimates and the actual reported service costs in all the years prior to the year whose service cost I am forecasting.

I begin my service cost forecasts in 1993 to allow changes in interest and compensation rates between 1991 and 1992 to be used for predicting 1993 service cost. 1993 is used because the rate information necessary is not available in Compustat until 1991 and a two-year lag is necessary to provide initial changes in rates for the estimation process. Based on the disclosed changes in the compensation and discount rates and the firm's 1992 service cost, I predict each firm's 1993 service cost by cycling through various n, m, and shift estimator combinations. I let n (m) vary between 0 and 20 (15), in half-percent intervals, and tried three different duration decile shift values (0 percent, 0.5 percent, and 0.75 percent), allowing n and m to take on independent shift estimators.[16] For each combination of n, m, and associated shifts tried, I made firm-specific adjustments in n and m values based on the firm's duration decile. Firms in the middle two duration deciles (deciles 5 and 6) receive no adjustment. The values of n and m are adjusted upward for firms in deciles below 5 (longer duration firms) and downward for firms in deciles above 6 (shorter duration firms) as follows. Firms in deciles 3 and 4 (7 and 8) get an adjustment of $+1x$ ($-1x$), where x is the shift estimator being tried. In deciles 2 and 9, the adjustment is $\pm 2x$, and in deciles 1 and 10 it is $\pm 3x$.

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[13] To ensure that the interest cost ratio is a good proxy for firms missing ABO, I compute the Pearson and Spearman correlation coefficients between the two proxies and find them both to be .52.

[14] I also run the estimations using duration deciles based on firm age. The results using the age proxy are similar.

[15] I also attempted to predict the future service cost using a more detailed set of equations based on assumed durations. These models proved more complex, relied on more intricate assumptions, and proved no more accurate.

[16] A total of 10,800 ($40 \times 3 \times 30 \times 3$) different n, m, and associated shift combinations.

As an example, consider the following combination: $n = 15$; n-decile shift estimator $= 0.75$; $m = 5$; m-decile shift estimator $= 0.5$. For a firm in duration decile 5 or 6, the 1993 service cost forecast (FSC1993) would be computed as:

FSC1993 = (Service Cost for 1992) * [1 − (Percent Change in Discount Rate) * (.15 + 0)] * [1 + (Percent Change in Compensation Rate) * (.05 + 0)] * (1 + Discount Rate).[17]

However, for a firm in, say, decile 1, this combination would yield a 1993 service cost forecast equal to:

FSC1993 = (Service Cost for 1992) * [1 − (Percent Change in Discount Rate) * (.15 + 3 * .0075)] * [1 + (Percent Change in Compensation Rate) * (.05 + 3 * .005)] * (1 + Discount Rate).

Note that for the discount rate, the n percent change term enters with a negative sign. This is because an increase in the discount rate results in a decrease in the service cost.

The percentages and associated shifts that produce the smallest error in service cost prediction when compared to the actual service cost in 1993 are then used to forecast the 1994 service cost. This algorithm is repeated for each successive year using all past firm years for the estimation. This methodology, therefore, yields slightly different n values, m values, and shift estimators for each year. On the whole, however, the values stay fairly constant with a 1 percent change in discount rate causing around a 14 percent change in service cost for firms in the middle duration deciles with a ±.75 percent shift per duration decile group change. A 1 percent change in compensation rate results in approximately a 5.5 percent change in service cost with a shift of 0 percent or ±.75 percent per duration decile group, depending on the year. Note that more accurate estimation methods might be possible if forecasts were done on an individual firm-by-firm basis rather than in the aggregate.

APPENDIX B

Pension Parameters and Control Variables Used in the Paper

Pension parameters used throughout the paper are defined as follows:

Total Assets = Compustat #6;
Plan Assets = Compustat (#287 + #296);
Plan Benefit Obligation (PBO) = Compustat (#286 + #294);
Prior Service Cost = Compustat (#288 + #288);
Unrecognized Gains and Losses + Transition Liability = Compustat (#289 + #299);
Service Cost = Compustat #331;
Interest Cost= Compustat #332;
Expected Return on Plan Assets (dollar value) = Compustat #333;

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[17] In the absence of any change in the discount rate the service cost should naturally increase due to the passage of time because retirees are now one year closer to their payouts. The (1 + Discount Rate) term in the estimation equations account for this.

Funded Status = (Plan Assets − Plan Benefit Obligation);
Off-Balance-Sheet Liabilities = (Prior Service Cost + Unrecognized Gains and Losses + Transition Liability);
On-Balance-Sheet Liabilities = −(Funded Status + Off-Balance-Sheet Liabilities);
Expected Rate of Return on Plan Assets (ERR) = Compustat #336;
Discount Rate = Compustat #246; and
Compensation Rate = Compustat #335.

Control variables used throughout the paper are defined as follows:

Market Value = Shares Outstanding * Fiscal Year-End Closing Price = Compustat (#25 * #199);
Book-to-Market = Book Value (Compustat #216)/Market Value;
Earnings-to-Price = Fiscal Year Closing Stock Price/EPS = Compustat (#199/#58);
Total Accruals = [Earnings before Extraordinary Items − (Cash Flow − Extraordinary Items)]/Lagged TA;
= Compustat (#123 − (#308 − #124))/#6;
Net Operating Assets (NOA) = (Operating Assets − Operating Liabilities)/Lagged Total Assets;
= Compustat [(#6 − #1) − (#6 − #34 − #9 − #38 − #130 − #60)]/#6;
Momentum = cumulative market-adjusted return for the six-month period ending one month before the start date; returns are obtained from CRSP;
Beta = beta deciles are obtained from the CRSP database;
Size = size deciles are obtained from the CRSP database; firms are ranked based on their market value relative to cutoffs established based on a decile ranking of NYSE firms in December of the previous year;
Analyst Forecast Error = the difference in dollars per share between actual firm earnings and the analyst consensus forecast constructed using I/B/E/S data; and
Age = the difference between a firm's fiscal year and its first recorded fiscal year in CRSP.

REFERENCES

Abarbanell, J., and V. Bernard. 1992. Tests of analysts' overreaction/underreaction to earnings information as an explanation for anomalous stock price behavior. *The Journal of Finance* 47 (3): 1181–1208.

———, and B. Bushee. 1997. Fundamental analysis, future earnings, and stock price. *Journal of Accounting Research* 35 (1): 1–24.

———, and ———. 1998. Abnormal returns to a fundamental analysis strategy. *The Accounting Review* 73 (1): 19–45.

Barth, M. E. 1991. Relative measurement errors among alternative pension asset and liability measures. *The Accounting Review* 66 (3): 433–463.

———, W. H. Beaver, and W. R. Landsman. 1992. The market valuation implications of net periodic pension cost components. *Journal of Accounting and Economics* 15: 27–62.

Bernard, V., and J. Thomas. 1990. Evidence that stock prices do not fully reflect the implications of current earnings for future earnings. *Journal of Accounting and Economics* 13 (4): 305–340.

Biddle, G. C., and W. E. Ricks. 1988. Analyst forecast errors and stock price behavior near the earnings announcement dates of LIFO adopters. *Journal of Accounting Research* 26 (2): 169–194.

Bradshaw, M., S. Richardson and R. Sloan. 2001. Do analysts and auditors use information in accruals? *Journal of Accounting Research* 39 (1): 45–73.

Brown, S. 2002. The impact of pension assumptions on firm value. Working paper, Emory University.

Chen, K., and M. Schoderbek. 2000. The 1993 tax rates increase and deferred tax adjustments: A test of functional fixation. *Journal of Accounting Research* 38 (1): 23–45.

Coronado, J. L., and S. A. Sharpe. 2003. Did pension plan accounting contribute to a stock market bubble? Brookings Papers on Economic Activity 1. Washington, D.C.: Brookings Institution Press.

Davis-Friday, P. Y., L. B. Folami, C. Liu, and H. F. Mittelstaedt. 1999. The value relevance of financial statement recognition vs. disclosure: evidence from SFAS No. 106. *The Accounting Review* 74 (4): 403–423.

———, J. Miller, and F. Mittelstaedt. 2005. Market-related values and pension accounting. Working paper, University of Notre Dame.

Elgers, P. T., and M. H. Lo. 1994. Reductions in analysts' annual earnings forecast errors using information in prior earnings and security returns. *Journal of Accounting Research* 32 (2): 290–303.

Financial Accounting Standards Board (FASB). 1985. *Employers' Accounting for Pensions*. Statement of Financial Accounting Standard No. 87. Stamford, CT: FASB.

———. 1998. *Employers' Disclosures about Pensions and Other Post-Retirement Benefits—An Amendment of FASB Statements 87, 88, and 106.* Statement of Financial Accounting Standard No. 132. Norwalk, CT: FASB.

———. 2003. *Employers' Disclosures about Pensions and Other Post-Retirement Benefits—An Amendment of FASB Statements No. 87, 88, and 106 and a Replacement of FASB Statement No. 132.* Proposed Statement of Financial Accounting Standards. Norwalk, CT: FASB.

Gold, J. 2000. Accounting/actuarial bias enables equity investment by defined benefit pension plans. Pension Research Council Working Paper 2001-5. Philadelphia, PA: The Wharton School, University of Pennsylvania.

Hirst, D. E., and P. E. Hopkins. 1998. Comprehensive income reporting and analysts' valuation judgments. *Journal of Accounting Research* 56 (Supplement)

Mikhail, M. B., B. R. Walther, and R. H. Willis. 1999. Does forecast accuracy matter to security analysts? *The Accounting Review* 74 (2): 185–200.

Plumlee, M. A. 2003. The effect of information complexity on analysts' use of that information. *The Accounting Review* 78 (1): 275–296.

Rajgopal, S., T. Shevlin, and M. Venkatachalam. 2003. Does the stock market fully appreciate the implication of leading indicators for future earnings? Evidence from order backlog. *Review of Accounting Studies* 8: 461–492.

Editorial Data

The following table contains information about turnaround time for manuscripts (including revisions) on which editorial decisions were made in the 12-month period ended December 31, 2005. Turnaround time is the number of days between the date that the manuscript was received *and the date of the editor's letter to the author(s).*

	Number of Manuscripts	Percent	Cumulative Number	Cumulative Percent
0 ≤ Days ≤ 30	84	18.06	84	18.06
31 ≤ Days ≤ 60	183	39.36	267	57.42
61 ≤ Days ≤ 90	140	30.11	407	87.53
91 ≤ Days ≤ 120	49	10.54	456	98.07
121 ≤ Days	09	1.93	465	100.00
Total	465	100.00		

The mean review time was 55.93; the median review time was 77 days.

New Submissions by calendar year (January 1 to December 31)

1992	241
1993	234
1994	231
1995	195
1996	230
1997	195
1998	196
1999	239
2000	260
2001	260*
2002	324
2003	327
2004	307
2005	389

* 2001—in addition to the 260 regular submissions, 68 MSs were submitted to the *TAR* Quality of Earnings Conference.

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EDITORIAL POLICY AND STYLE INFORMATION

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Dye, R., B. Balachandran, and R. Magee. 1989. Contingent fees for audit firms. Working paper, Northwestern University, Evanston, IL.
Fabozzi, F., and I. Pollack, eds. 1987. *The Handbook of Fixed Income Securities*. 2nd edition. Homewood, IL: Dow Jones-Irwin.
Kahneman, D., P. Slovic, and A. Tversky, eds. 1982. *Judgment Under Uncertainty: Heuristics and Biases*. Cambridge, U.K.: Cambridge University Press.
Porcano, T. M. 1984a. Distributive justice and tax policy. *The Accounting Review* 59 (October): 619–636.
———. 1984b. The perceived effects of tax policy on corporate investment intentions. *The Journal of the American Taxation Association* 6 (Fall): 7–19.
Shaw, W. H. 1985. Empirical evidence on the market impact of the safe harbor leasing law. Ph.D. dissertation, The University of Texas at Austin.
Sherman, T. M., ed. 1984. *Conceptual Framework for Financial Accounting*. Cambridge, MA: Harvard Business School.

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402 UNIVERSITY OF IDAHO, Department of Accounting seeks applicants for tenure-track faculty appointment as Assistant or Associate Professor of accounting beginning July 2006. Teaching emphasis in managerial, information systems, and financial accounting. Salary depends on qualifications and experience. Ph.D. or D.B.A. in accounting and demonstrated interest and/or record of teaching, research, and service accomplishments required. Professional certification desired. ABDs will be considered. Applicants should apply online at: http://www.hr.uidaho.edu (job seekers) or send applications including a complete resume, a graduate studies transcript, teaching evaluations, and letters from three professional references to: Marla Kraut, Department of Accounting, University of Idaho, PO Box 443161, Moscow, ID 83844-3161; Phone: (208) 885-6453. Screening will begin immediately. AA/EOE.

403 CHINA EUROPE INTERNATIONAL BUSINESS SCHOOL, (CEIBS) invites applications for the position of Associate Professor of accounting, beginning September 2006 or January 2007. Candidates must have a Ph.D. in accounting, outstanding teaching record for M.B.A., E.M.B.A. and/or Executive Education in English, good understanding of management practice (emphasis on China a plus) and a proven research record. CEIBS is an independent business school located in Shanghai, China. The school's M.B.A. Program was ranked 21st by *Financial Times* in February 2006; E.M.B.A was ranked 13th by *FT* last October; Open Program of Executive Education was ranked 37th by *FT* last May. More information can be obtained at: http://www.ceibs.edu. Interested candidates should send an application letter, *curriculum vitae*, names and addresses of three references, plus any supporting documents by August 15, 2006 to: Helen Xu, Dean's Office, CEIBS, 699 Hongfeng Road, Shanghai, 201206, PRC. Email submissions are encouraged: xhuijuan@ceibs.edu.

404 UNIVERSITY OF NOTRE DAME, Department of Accountancy, invites applications for position openings at the Assistant and Associate Professor levels. All research areas will be considered. Applicants for Assistant Professor positions should expect to complete doctoral degree requirements before the start of the contract period. All candidates should have a strong commitment to scholarly research and excellence in teaching. Salary is very competitive and research support is excellent. Please send *curriculum vitae* and an example of scholarly work to: Professor Tom Schaefer, Chair, Department of Accountancy, Room 102D, Mendoza College of Business Administration, University of Notre Dame, Notre Dame, IN 46556-5646. The University of Notre Dame is an Equal Opportunity/Affirmative Action Employer.

405 NAVAL POSTGRADUATE SCHOOL, Graduate School of Business and Public Policy (AACSB- and NASPAA-accredited) seeks all ranks of tenure-track faculty in Accounting and Finance. Applicants must possess good research skills, a doctoral degree in accounting or finance (or close to completion), and evidence of strong commitment to M.B.A. teaching. U.S. citizenship is required for award of tenure. Visiting positions will also be considered. Applications by email preferred; please send covering letter, current resume, and three letters of recommendation or contact information for references to: busapp@nps.edu. Applicants should submit evidence of teaching abilities and scholarship. Attachments should be in PDF or Word (no zipped files please). Applications may also be sent to: Douglas Moses, Senior Associate Dean, Graduate School of Business and Public Policy, Mail Code GB/Mo, Naval Postgraduate School, 555 Dyer Road, Building 330, Ingersoll Hall, Room 307, Monterey, CA 93943. For more information, see http://www.nps.navy.mil/gsbpp/. NPS is an Equal Opportunity/Affirmative Action Employer.

406 Ph.D. in accounting (Big-Ten school) with thirty-plus-years of building one of the top ranked, non-doctorate-granting accounting programs. Have taught across the accounting curriculum except for auditing and tax. My primary niche is governmental and not-for-profit area. Have worked closely with GASB for more than ten years. Am interested in exploring new opportunities, either domestically or internationally. Available ?? Send any ideas to: govacctguy@yahoo.com.

407 UNIVERSITY OF MARYLAND UNIVERSITY COLLEGE, (UMUC) has an opportunity for an academic accounting professional to lead in an innovative and global environment and make a difference. Position: Collegiate Faculty/Academic Director, Accounting School of Undergraduate Studies; 12-month faculty; FT position with initial two-year contract. UMUC will fund up to set amount toward relocation. Reports to: Chair of Accounting, Economics and Finance. Responsibilities: Lead/direct 115 collegiate (13)/adjunct (102) faculty; Develop/manage overall Accounting program: faculty management/support, curriculum/program development, program assessment/review; Teach one course/year. Qualifications: Terminal degree (Ph.D. Accounting/related discipline); CPA; five years of teaching in four-year institution and five years of management experience; Outstanding communication skills and ability to interact internally/externally; Strong motivator/leader aware of accounting developments/practices in public/industry; Preferred: seven-plus years of managing academic center/distance education. Submit cover letter, résumé, salary history, and three references to: Beth Berk, CPA, Independent Recruiter. Email: JobsBAB@comcast.net, note "JobAAA001254" in subject line.

408 UNIVERSITY OF MARYLAND, seeks applicants for a continuing non-tenure-track teaching position in Accounting. The position is for nine months, beginning late January 2007 or late August 2007, and is renewable subject to performance. Off-term teaching is available. Candidates must possess a doctorate and an excellent teaching record, especially at the Master's level. Teaching breadth will be valued. Certain student support and service responsibilities will be expected. Send *curriculum vitae*, three letters of recommendation, and evidence of teaching effectiveness to: Professor Jim Peters, The Robert H. Smith School of Business, University of Maryland, College Park, MD 20742-1815. Women and minorities are especially encouraged to apply. The University of Maryland is an Equal Opportunity/Affirmative Action Employer.

409 LOUISIANA TECH UNIVERSITY, School of Professional Accountancy, seeks applications for Assistant or Associate Professor for Fall 2007 appointment. Accounting doctorate is required and certification preferred. An established publication record of discipline-based scholarship in high-quality journals is essential. Position requires teaching at the upper division and graduate levels, publishing discipline-based research, and involvement with the accounting doctoral program. Interest in auditing is preferred, but other interests will be considered. Applicant review will continue until position is filled. Salary is highly competitive. Located 65 miles east of Shreveport and 30 miles west of Monroe, enrollment totals approximately 12,000 students. Please send applications or letters of inquiry to: Thomas J. Phillips, Jr., Director, School of Professional Accountancy, PO Box 10318, Louisiana Tech University, Ruston, LA 71272. Phone: (318) 257-2822; Fax: (318) 257-4253; Email: phillips@cab.latech.edu; Website: http://www.cab.latech.edu. Women and minorities encouraged to apply. Member of the University of Louisiana System. EEO/AA/ADA.

410 UNIVERSITY OF DELAWARE, Alfred Lerner College invites applicants for a tenure-track Assistant Professor of Taxation, Accounting, and MIS position beginning September 2007. Primary qualifications include outstanding research potential, teaching ability, a doctorate in Accounting, or be ABD. Experienced Assistant Professors are encouraged to apply. Candidates are expected to teach undergraduate/graduate-level courses. Salary, benefits, and support are competitive with AACSB-accredited schools. Will interview applicants at the 2006 American Accounting Association meeting in Washington, D.C. Send three letters of reference, *curriculum vitae*, and teaching performance summaries to: Kent St. Pierre, Chair, Search Committee, Alfred Lerner College of Business and Economics, Department of Accounting and MIS, University of Delaware, Newark, DE 19716, 302-831-2961. May apply electronically to: warrenn@lerner.udel.edu. *Curriculum vitae* and letters of reference shall be shared with departmental faculty. The University of Delaware is an Equal Opportunity Employer, and encourages applications from Minority Group Members and Women.

411 ST. MARY'S UNIVERSITY, Bill Greehey School of Business invites applications to fill a tenure-track position in accounting at the Assistant/Associate/Full Professor level. Discipline expertise in auditing and information systems strongly desired. Expertise in other area(s) of accounting is a plus. Candidates are required to hold a Ph.D. or D.B.A. from an AACSB-accredited institution, and demonstrate the ability to bring excellence and innovation into the classroom, maintain a strong and continuing record of scholarship consistent with the teacher-scholar model, and possess attributes leading to successful professional service. Salary is competitive. The position begins by August 2007. Interested candidates should send an application letter, current *vita*, and names with contact information for three references to: Dr. Tom Madison, Chair, Department of Accounting, Bill Greehey School of Business, St. Mary's University, One Camino Santa Maria, San Antonio, TX 78228-8607. Deadline for receipt of materials is October 1, 2006.

412 WORCESTER POLYTECHNIC INSTITUTE, Department of Management seeks applications for a tenure-track opening in Managerial Accounting. The selected faculty member will teach three or four courses per year, advise student projects, publish in high-quality journals, and engage in appropriate service. The ideal applicant will hold a Ph.D. in accounting, and have teaching experience in managerial accounting, as well as a demonstrated ability to convert research into scholarly publications. WPI is a selective private university with an innovative project-enriched curriculum that engages students and faculty in real-world problem solving, often at corporate and international sites. Located in Worcester, Massachusetts, New England's third largest city and ranked 53rd by *US News & World Report* among national comprehensive universities, WPI enrolls 3,900 undergraduate and graduate students. For additional information, including how to apply, please visit our website at: http://www.wpi.edu/Admin/HR/Jobs/.

413 MERRIMACK COLLEGE, Department of Accounting invites applications for a tenure-track position at the Assistant or Associate Professor level beginning Fall 2006. The department is seeking an individual holding a doctoral degree in Accounting by the start of the appointment. Candidates must have a commitment to excellence in teaching, research, and professional service. This position includes teaching 12 credits per semester, advising as well as participation on committees. Please submit your letter of application and *curriculum vitae* to: Erica Lessinger, Office of the Dean, Merrimack College, 315 Turnpike Street, Box O5, North Andover, MA. 01845; or email: erica.lessinger@merrimack.edu; or apply online at: http://www.merrimack.edu/jobs. Women and minorities are encouraged to apply. Merrimack College is an Equal Opportunity Employer and has a strong institutional commitment to diversity.

414 THE GERMAN JORDANIAN UNIVERSITY, newly established Talal Abu-Ghazaleh College of Business in Amman, Jordan, invites applications at all levels for tenure-track or visiting appointments in accounting and also in other disciplines for the Fall 2006 semester. All candidates should possess Ph.D. degrees and have a strong commitment to scholarly research and strong teaching. Applicants for senior positions should have a proven record of research, publication, and teaching. The German Jordanian University is a new public university formed by the governments of Germany and Jordan. The Talal Abu-Ghazaleh College of Business offers an undergraduate program in accounting and management, and an M.B.A. program in International Accounting and Quality Management. All instruction is done in English. Applications should be addressed to Recruiter and sent via email to: bag@tagi.com. Websites: http://www.tagcb.edu.jo and http://www.gju.edu.jo.

415 OHIO WESLEYAN UNIVERSITY, Department of Economics invites applications for an Assistant Professor of Accounting. This is a tenure-track position, beginning August 2006. The successful candidate will teach undergraduate courses in Managerial Accounting, Cost Accounting and Accounting Systems, and other topics. Master's Degree, CPA, and professional experience required. Teaching experience preferred. Additional job requirements include advising students, participating in department and university activities, and research or professional activities. Salary mid-$40s with a comprehensive benefit package. Send resume, transcript, and three references to: Professor Joann Harvey, Economics Department, Ohio Wesleyan University, Delaware, OH 43015; Email: jpharvey@owu.edu. Review of applications will begin on 5/5/06 and will continue until filled. Ohio Wesleyan University is an Affirmative Action/Equal Opportunity Employer strongly committed to diversity within its community.

416 MARQUETTE UNIVERSITY, invites applications for two tenure-track positions, one in Financial Accounting and one in any area except Tax, beginning Fall 2007. Applicants should hold a doctoral degree or be in the final stages of dissertation. The first position will be filled at the rank of Assistant Professor and primary teaching responsibility will be in Financial Accounting at the undergraduate and graduate levels. The second position will be filled at either the rank of Assistant or Associate Professor with primary teaching responsibility at the undergraduate and graduate levels. The normal teaching load is 12 semester hours per year the first year and 15 semester hours thereafter. Salary and research support are competitive. The department has 11 full-time faculty. Marquette offers AACSB-accredited undergraduate and graduate degrees in Accounting and Business. Please send *curriculum vitae* and inquiries to: Michael D. Akers, Professor and Chair, Department of Accounting, PO Box 1881, College of Business Administration, Marquette University, Milwaukee, WI 53201-1881; Email: Michael.Akers@Marquette.edu. Marquette University is an Affirmative Action/Equal Opportunity Employer.

417 UNIVERSITY OF HONG KONG, invites applications for Professor/Associate Professor/Assistant Professor in Accounting in the Faculty of Business and Economics (Ref.: RF-2005/2006-411), from August 1, 2006 for three-year fixed-term with the possibility of renewal or on tenure for exceptionally outstanding candidates (except for Assistant Professor level). Applicants should have a Ph.D. degree in Accounting, be able to teach Auditing, Financial, and/or Managerial accounting subjects at the undergraduate and graduate levels, demonstrate excellence in teaching, commit to high-quality scholarly research, publish in top-tier academic journals, and supervise doctoral students. Professional qualification is also preferred. Candidates for the rank of Associate Professor/Professor should have a proven record of high-quality scholarship, teaching and services, and be expected to conduct an ongoing active research agenda, obtain research grants, and serve on committees within the University and external professional organizations. Further details of the post and application procedure are available at: https://extranet.hku.hk/apptunit/. Review of applications will start from May 2006.

418 FULLER & THALER ASSET MANAGEMENT, INC., invites applications for a Quantitative Investment Analyst. Fuller & Thaler is a rapidly growing investment management firm based in San Mateo, California. Due to continued growth, Fuller & Thaler is seeking to expand its Quantitative Research group. Reporting to the Head of Quantitative Research, the selected person's responsibilities include conducting exploratory data analysis, developing and enhancing models, and reviewing financial and accounting literature. The ideal candidate will have the following qualities: Ph.D. in accounting or finance; Strong written and oral communication skills; Familiar with financial statement analysis; Knowledge of the accounting and financial literature; Experience with database programs such as SQL Server; Experience with both historical and real-time financial databases such as Compustat, CRSP, I/B/E/S, MSCI, Worldscope, Bloomberg, FACTSET, etc.; Experience with analytical programs such as Matlab and SAS. For confidential consideration for this position, please email your resume to: hr@fullerthaler.com or fax to Human Resources at: (650) 227-0136.

419 BENTLEY COLLEGE, Department of Accountancy seeks a professor with teaching and research interests in the area of Accounting Information Systems. We are looking for an individual who cares about students, conducts research that is relevant and meaningful to the business community, desires to work with some of the finest scholars in the accounting academic community, can help the department deliver a first-class Ph.D. in Accountancy, and contributes to the positive collegial atmosphere that exists among faculty members. The academic rank is open and the salary will be competitive. Please send letter of interest and *curriculum vita* directly to: Professor Mark Nixon, Chair, Department of Accountancy Bentley College, 175 Forest Street Waltham, MA 02452; Email: mnixon@bentley.edu; Phone: (78) 891-2087.

420 UNIVERSITY OF SOUTHERN MAINE, Department of Accounting and Finance invites applications for a tenure-track position in accounting beginning September 2007 with teaching and scholarship interests in auditing and upper-level financial accounting. Applicants must have an earned doctorate with a concentration in accounting from an institution accredited by AACSB International. ABDs who are near completion will be considered, but must have completed all degree requirements by the time of the appointment. This candidate is expected to be an effective teacher, publish quality research, and contribute to service activities. Professional experience and CPA are preferred. Women and minorities are encouraged to apply. Send a letter of application and *curriculum vitae* to: Dr. Jeff Shields, University of Southern Maine, School of Business, PO Box 9300, Portland, ME 04104-9300; Email: jshields@usm.maine.edu. Candidates will be expected to provide three letters of recommendation. Review of applications begins immediately.

421 UNIVERSITY OF AKRON, invites applications for two tenure-track positions (Auditing/Financial and Taxation) at the Assistant or Associate Professor rank, beginning January or August 2007. Required qualifications include a Ph.D., evidence of effective teaching, and a commitment to research and scholarship. For the tax position, we will consider applicants with combined J.D./Master's/CPA. The School of Accountancy, which has separate AACSB accreditation, offers B.S. (Accounting), M.S. (Accounting), and M.Tax degrees. We will interview applicants at the 2006 American Accounting Association meeting in Washington, D.C. Send a cover letter, names of three references, and *vita* to: Professor Thomas G. Calderon, Chair, George W. Daverio School of Accountancy, College of Business Administration, The University of Akron, Akron, OH 44325-4802; Phone: (330) 972-6228; Email: tcalderon@uakron.edu. Review of applications will be ongoing until the position is filled.

422 UNIVERSITY OF LOUISVILLE, School of Accountancy in the College of Business invites applications for a full-time, tenure-track position at the level of Assistant or Associate Professor to begin in August 2007. The successful candidate will have a terminal degree and strong accomplishment in both scholarly research and teaching in the area of auditing/assurance services. Preference will be given to candidates with professional designation(s). The Search Committee will accept applications until a new faculty member is selected. For optimal consideration, individuals should attach a letter of interest, *curriculum vita*, and two samples of scholarly work prior to August 1. Three current letters of recommendation should be mailed to: Dr. Richard Coppage, Accountancy Search Committee, University of Louisville, Louisville, KY 40292, or electronically to: coppage@louisville.edu. All candidates must apply online at: www.louisville.edu/jobs, Job Opening ID #15376. The University of Louisville is an Affirmative Action/Equal Opportunity/Americans with Disabilities.

423 LOYOLA MARYMOUNT UNIVERSITY, Department of Accounting invites applicants for two full-time tenure-track faculty positions at Assistant/Associate rank, effective Fall 2007. A doctorate (or short-term ABD status) in accounting from an AACSB-accredited university is required. All candidates should demonstrate a strong commitment to scholarly research and excellence in teaching. Our priority is for teaching in the AIS and Financial Accounting areas. Salary is competitive and research support is available. If you are interested in learning more about these positions, contact: Professor Alan Cherry, Department of Accounting, College of Business Administration, Loyola Marymount University, One LMU Drive, Los Angeles, CA 90045-2659. Loyola Marymount, a comprehensive university in the mainstream of American Catholic higher education, seeks professionally outstanding applicants who value its mission and share its commitment to academic excellence, the education of the whole person, and the building of a just society. Women and minorities are encouraged to apply. LMU is an Equal Opportunity institution, actively working to promote an intercultural learning community. Visit http://www.lmu.edu for more information.

424 THE UNIVERSITY OF TEXAS AT EL PASO, Department of Accounting invites applications for a tenure-track position, specialization in Taxation, Accounting Systems, or Auditing. Rank is open. Candidates must have a Ph.D. or other terminal degree in accounting prior to appointment. Applicants for rank above Assistant Professor must have demonstrated accomplishment in publication, teaching, and service. Professional certification is highly desirable. UTEP's College of Business Administration offers degrees at the Bachelor's and Master's levels, and a Ph.D. in International Business. The Department and College are both accredited by AACSB and salary is competitive. Review of applications will begin immediately and continue until the position is filled. Send application, *curriculum vitae*, and names of five references to: Dr. Pat Eason, Chair, UTEP Department of Accounting, El Paso, Texas 79968. UTEP does not discriminate on the basis of race, color, national origin, sex, religion, age, disability, veteran status, or sexual orientation.

425 UNIVERSITY OF COLORADO AT DENVER, Business School and Health Sciences Center seeks applicants for accounting tenure-track faculty. Minimum qualifications for Assistant Professor include a Ph.D. in accounting and a commitment to excellence in research and teaching. Candidates for advanced rank (Associate or Full Professor) must demonstrate significant accomplishments in research, teaching, and service. Specialization in managerial accounting or accounting information systems is preferred, though all areas will be considered. Applicants who demonstrate proficiency using new accounting or information technologies in their teaching are especially encouraged to apply. Review of applications begins immediately, and will continue until filled. Appointment can begin Fall Term 2007, or as negotiated, and is dependent on funding availability. Applicants for Visiting Professor are also encouraged to apply. The Carnegie Foundation has just classified UCDHSC as RU/VH: Research Universities (very high research activity). For full description, please see: http://www.ucdhsc.edu/admin/hr/facpos.htm#asstprof The Business School website is: http://www.cudenver.edu/Academics/Colleges/Business/default.htm. Contact: Professor Bruce R. Neumann, Director of Accounting Programs, PO Box 173364, Campus Box 165, Denver, CO 80217-3364; Email: Bruce.Neumann@cudenver.edu.

426 CANISIUS COLLEGE, Wehle School of Business, Buffalo, New York, invites applicants for one tenure-track position in Accounting. Applicants for the position must possess a Ph.D. with a concentration in accounting or be near completion of the dissertation. Applicant's primary interest should be Advanced Accounting and Theory. The Department of Accounting seeks faculty with a strong commitment to teaching excellence balanced with an active interest in scholarly research to join its nationally recognized program. Canisius offers both graduate and undergraduate accounting degrees. The Wehle School is accredited by the AACSB International. Send letters of interest and *curriculum vitae* to: Dr. Ian J. Redpath, Department of Accounting, Wehle School of Business, Canisius College, 2001 Main St., Buffalo, NY 14208-1098. Canisius College, a Jesuit College, is an Equal Opportunity Employer committed to recruiting, supporting, and fostering a diverse community of outstanding faculty, staff, and students. Visit: http://www.canisius.edu/accjob.

427 SANTA CLARA UNIVERSITY, Department of Accounting is authorized to hire two tenure-track professors, at any level, for Fall 2007. All teaching and research areas will be considered, but priority will be given to financial accounting, managerial accounting, and accounting information systems. Candidates for Assistant Professor should have completed or nearly completed their Ph.D. Candidates at higher ranks should have a record of research publications in top-tier journals and evidence of teaching excellence. Address inquiries and correspondence to: Dr. Michael Eames, Department of Accounting, Leavey School of Business, Santa Clara University, Santa Clara, CA 95053-0380. Email to: meames@scu.edu; Phone: (408) 554-4896). SCU is a private Jesuit university located in Silicon Valley with competitive salaries, housing subsidies, and summer research support. SCU is an EEO/AA employer.

428 CLAREMONT MCKENNA COLLEGE, Economics/Accounting Department invites nominations/applications for a tenure-track position in Accounting beginning Fall 2007, subject to budgetary approval. We seek an individual holding a Ph.D. in Accounting, with a commitment to excellence in teaching, research, and professional service. The position offers competitive compensation and teaching responsibilities that enable a notable level of high-impact research. Rank and salary dependent on qualifications. CMC is an elite residential liberal arts college near Los Angeles. CMC specializes in economics and public affairs, with an Economics/Accounting faculty of 20, including five Accounting faculty members. http://econ.claremontmckenna.edu/majors/econ-acct/. Applications will be reviewed immediately. Position will remain open until filled. A representative will attend the AAA meeting. Submit cover letter, resume, teaching evaluations, references, and publications/working papers to: Chair, Accounting Search, Economics/Accounting Department, Claremont McKenna College, 500 East Ninth Street, Claremont, CA 91711. Claremont McKenna College is an Equal Opportunity Employer.

429 LEBANON VALLEY COLLEGE, invites applications for a full-time position beginning in August 2007. Candidates must have a current CPA license along with an earned M.B.A. or a Master's degree or Doctorate in accounting. Candidates must have a record of teaching excellence and be committed to learning outcomes assessment. Faculty are expected to teach, advise majors, participate in scholarship and/or professional development, and serve on faculty committees. Deadline for applications is September 15, 2006. Send a complete application, including cover letter, statements of teaching philosophy and research/development agenda, *c.v.*, and reference list to: Dr. David V. Rudd, Department of Business and Economics, 101 N. College Avenue, Annville, PA 17003. Salary and rank are commensurate with qualifications. For additional information on this position see: http://www.lvc.edu. LVC is an Equal Opportunity Employer and encourages applications from women and under-represented minorities.

430 NORTHERN ILLINOIS UNIVERSITY, invites applications for anticipated tenure-track openings for accountancy faculty at the rank of Assistant or Associate Professor beginning Fall 2007. We are primarily interested in candidates who have a teaching interest in Financial Accounting, International Accounting, or Auditing. NIU offers B.S. in Accountancy, Master of Accounting Science, and Master of Science in Taxation degrees. Information about our nationally ranked programs can be found at: http://www.cob.niu.edu/accy/. Candidates must be qualified to teach at the undergraduate and graduate levels. Candidates should have, or expect to have, a Ph.D. degree when employment begins. Teaching experience, business experience, and professional certification are desirable. We are particularly interested in faculty who want to combine their research and teaching with a close working relationship with the professional business community. NIU representatives will be available at the AAA Meeting in Washington, D.C. Applicants must send a cover letter and *curriculum vitae* to: Jim Young, Chair, Department of Accountancy, Northern Illinois University, DeKalb, IL 60115. Complete applications must be received by August 18, 2006. AA/EEO.

ILLINOIS STATE UNIVERSITY
Normal/Bloomington

Faculty Position – Accounting
Contingent on Budgetary Approval and Appropriate Funding

The Department of Accounting at Illinois State University invites applications for a full-time, tenure-track Assistant or Associate Professor position in the Department of Accounting, preferably to begin August 16, 2007. Qualifications: Ph.D. or DBA in Accounting required, with a strong commitment to quality teaching, research, and service. ABD's will be considered if substantial progress has been made on the dissertation. Preference will be given to applicants with a primary teaching interest in accounting information systems and a secondary teaching interest in management or financial accounting. To assure full consideration, please send a letter of application and evidence of qualifications by September 1, 2006 to: Gerry McKean, Interim Chairperson, Department of Accounting, Campus Box 5520, Illinois State University, Normal, IL 61790-5520, e-mail: gerry.mckean@ilstu.edu.

Illinois State is an equal opportunity/affirmative action university encouraging diversity

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ILLINOIS STATE UNIVERSITY
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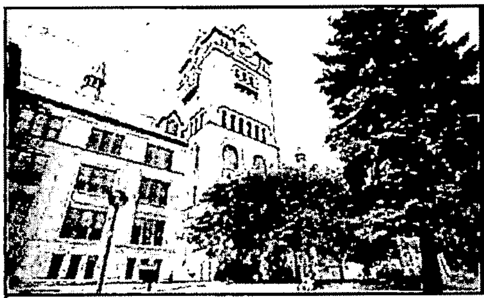

WAYNE STATE UNIVERSITY

SCHOOL OF BUSINESS ADMINISTRATION

DEAN

Wayne State University is one of the country's leading urban doctoral-extensive universities, providing access to a world-class education at a great value. Wayne State's 11 schools and colleges offer more than 350 major subject areas to over 33,000 graduate and undergraduate students. Located in the heart of Detroit's University Cultural Center, the home of renowned museums, galleries and theatres, the WSU main campus encompasses 203 acres of beautifully landscaped walkways and gathering spots, linking 100 education and research buildings. A number of new residential communities and a new 75-acres research and technology park contribute to the vitality of campus life.

With 58 full-time faculty, the AACSB International Accredited School of Business Administration serves nearly 2500 undergraduate and over 1200 graduate students pursuing degrees in the fields of accounting, finance, information systems & manufacturing, logistics, management, and marketing. Currently, the School is in the final stages of approval of a doctoral program. The MBA program, which includes one of the nation's largest part-time programs is supported by an award-winning online education initiative. Preparing business leaders for more than 60 years, the School's 30,000 alumni include many who hold key executive positions in major corporations, nonprofit organizations, and government agencies.

The University is seeking a Dean who can lead the School in a mission of academic excellence. The Dean should be an experienced academic administrator or business executive who can lead the faculty and stakeholders in furthering the undergraduate and graduate business degree programs and faculty research and development. The Dean is responsible for advancing the School's strategy and leveraging the School's resources to build its brand identity, alumni loyalty, and external relations. The Dean will be expected to identify and build on opportunities to work with business firms located throughout the state of Michigan and around the globe. Wayne State University, in the midst of its first-ever capital campaign, is raising $500 million for university-wide initiatives that will transform the learning and teaching experience on campus. The Dean will work with University officials, alumni, leaders of Fortune 500 companies, and prominent regional business firms to raise the $45 million targeted for construction of a new building for the School, and endowments for student scholarships and chaired professorships. The Dean should have a demonstrated record of accomplishment in a collegial leadership position and possess excellent interpersonal and communication skills. Experience with AACSB International accreditation is preferred.

The Dean reports to the Provost and Senior Vice President for Academic Affairs. The salary will be competitive and commensurate with experience and qualifications. Review of applications will begin immediately and continue until the position is filled. Applications must include a letter of intent, curriculum vitae, the names, addresses, and telephone numbers of three references, any evidence of recent scholarly activity, and evidence of other accomplishments. Electronic applications must be sent to **https://jobs.wayne.edu** Inquiries should be directed to: **Dean Ralph Kummler, Search Chair, Wayne State University, College of Engineering, 1150 Engineering Building, Detroit, MI 48202; Phone: 313-577-3775; Fax: 313-577-5300; email: rkummler@eng.wayne.edu**

Please refer to posting number 032770.

Wayne State University is an Equal Opportunity/Affirmative Action Employer.

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Accounting Quality and Firm-Level Capital Investment

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ABSTRACT: This study examines how accounting quality relates to firm-level capital investment efficiency. Our first hypothesis is that higher quality accounting enhances investment efficiency by reducing information asymmetry between managers and outside suppliers of capital. Our second hypothesis is that this effect should be stronger in economies where financing is largely provided through arm's-length transactions compared with countries where creditors supply more capital. Our results are consistent with these hypotheses both across and within countries. They are robust to alternative econometric specifications, different measures of accounting quality and investment-cash flow sensitivity, and numerous control variables.

Keywords: *accounting quality; capital investment; corporate governance.*

I. INTRODUCTION

A significant body of prior literature examines relations between accounting quality and financial market characteristics, yet little prior research exists that analyzes the effects of accounting quality on investment. Here, we examine how accounting quality relates to firm-level capital investment efficiency, a key determinant of economic productivity. Our focus on investment in productive assets complements and extends findings on how accounting relates to public financial markets' operating characteristics. For example, Bhattacharya et al. (2003) find that accounting opacity is associated with a higher cost of publicly traded equity capital across 34 countries. However, institutional features may be related to firm-level capital investment differently than to equity market operations. For instance, private financing from banks in Germany (e.g., Gorton and Schmid 2000), *keiretsu* in Japan (e.g., Hoshi et al. 1991), and families in East Asia (e.g., Claessens et al. 2000) may obviate or alter the effects of institutional features from their operational roles

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in public equity markets. Hence, how accounting quality affects the efficiency of firm-level capital investments remains an open question of fundamental importance.

In the neoclassical setting, managers (i.e., firms) endowed with capital invest until the marginal return is zero, allowing for adjustment costs (e.g., Tobin 1969; Hayashi 1982). In this baseline setting, we should not observe an association between internally generated cash flows and investment. But several frictions contradict this efficient result. One is capital rationing, which is generated by information asymmetry between managers and investors. Since at least Myers (1984), it has been shown that if managers can exploit private information to issue securities at inflated prices, then investors rationally withhold capital. A resulting reliance on internal funding increases the sensitivity of investments to cash. A second friction arises from agency problems, when managers pursue perquisite consumption and "empire building" rather than returning excess cash to investors (e.g., Jensen 1986; Blanchard et al. 1994).[1] This behavior also may increase the sensitivity of investments to cash flows, but, in this case, the sensitivity is due to an excess of cash.

Certain institutional features may serve to mitigate these deviations from the optimal investment policy. For example, Rajan and Zingales (2000) observe that "to function properly, a financial system requires clear laws and rapid enforcement, an accounting and disclosure system that promotes transparency, and a regulatory infrastructure that protects consumers and controls risk." Transparent accounting should reduce both adverse selection (i.e., the tendency to issue securities at an inflated price) and moral hazard (i.e., perquisite consumption using assets in place) by improving contracting and monitoring.[2] Thus, higher quality accounting may serve to enhance investment efficiency by mitigating these frictions.

We examine this hypothesis by considering the effects of differences in accounting quality both at the country level and at the firm level within countries. To do so, we first estimate investment-cash flow sensitivities for firms from 34 countries. We then consider how the average investment-cash flow sensitivity across countries varies with accounting quality. We measure accounting quality using three proxies for earnings quality described by Bhattacharya et al. (2003), as well as a measure of accounting timeliness from Bushman et al. (2004). We also are careful to control for the confounding effects of other institutional features, such as disclosure quality (CIFAR index), legal origin, creditor and shareholder rights, judicial efficiency, and economic conditions (e.g., the degree of economic development). We find that higher accounting quality is associated with lower investment-cash flow sensitivity. We also find that other institutional features, particularly creditor rights and disclosure quality, play a similar but incremental role. These results persist after allowing for the possibility that operating cash flows convey additional information about short-term profitability (e.g., Alti 2003).

Having established that differences in accounting quality are associated with differences in the efficiency of capital investments across countries, we then examine how sources of financing (debt versus equity) affect this relation. *A priori*, we expect accounting quality to play a stronger role in economies where financing is largely provided through arm's-length transactions, for example, where stock exchanges are the dominant sources of capital, since here, investment decisions rely more heavily on public accounting disclosures. On the other hand, in economies where creditors play a more dominant role, banks may be able to obtain information through private channels (mitigating adverse selection problems), and they may be in a better position to directly monitor managers once capital is supplied

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[1] Note that the existence of the agency problem *ex post* may lead to rationing of capital *ex ante*.

[2] See Healy and Palepu (2001) for a review.

(mitigating moral hazard). Thus, in credit-dominated economies, the quality of publicly disclosed accounting information should be less influential in decisions to supply capital than in public equity-dominated economies. Therefore, we should observe a smaller effect on capital investment decisions when accounting quality is improved in credit-dominated economies than in public equity-dominated economies.

We test this hypothesis in two ways. Initially, we regress cash flow sensitivity on accounting quality in our cross-country sample, partitioning observations in two groups based on the prevalence of public equity versus debt financing. We find that higher accounting quality is associated with lower investment-cash flow sensitivity in economies that depend more on public equity financing compared with those more reliant on debt financing. Then, we examine the effect of accounting quality on investment efficiency at the firm level in two selected countries. We do this for several reasons. First, it allows us to determine whether accounting quality operates similarly across and within countries. Second, it helps us to control for certain differences in characteristics across countries that could affect our cross-country results, such as correlations between accounting quality, legal origins, creditor rights, and administrative efficiency. Finally, as explained below, this firm-level analysis allows us to address certain econometric issues.

In our within-country tests, we examine two polar cases—the U.S. and Japan—for several reasons. First, they are the two largest economies in the world with abundant external financing available to firms. Second, they provide large and diversified firm samples with long time-series data that enable us to estimate firm-specific parameters with a reasonable level of confidence. Comparable datasets are generally not available for other countries. However, the U.S. and Japan differ along one important dimension. Public equity capital plays a much more dominant role as a source of firm financing in the U.S. than in Japan. Thus, we expect accounting quality to play a more important role in the largely arm's-length transactions in the U.S. On the other hand, Japanese suppliers of capital, such as *keiretsu* and banks, have non-public sources of information and thus the quality of public accounting disclosures may be less relevant to their decisions to supply capital. Our results are consistent with these predictions. Accounting quality matters both statistically and economically in the U.S. where higher quality accounting is associated with lower investment-cash flow sensitivity, but we do not observe such a relation in Japan.

Our findings contribute in at least two ways to the existing literature. First, they provide empirical evidence that accounting and other institutional features relate to the economic fundamentals of firm-level capital investment efficiency. In particular, accounting quality is shown to reduce frictions in the investment process. We find that this effect exists both across countries and within countries, even in the most liquid capital market, the U.S. Second, our findings confirm that this effect is stronger in economies where public equity capital plays a greater role in capital investment financing compared with countries dominated by debt financing. This effect is observed in both our cross- and within-country tests.

We proceed as follows. Section II reviews related research. Section III develops our hypotheses. Section IV describes our tests for relations between accounting quality and investment-cash flow sensitivity. Section V presents our findings. Section VI concludes the paper.

II. RELATED RESEARCH

Economists have long studied how financial frictions affect investment decisions and economic growth. This literature is too extensive to review comprehensively here, though we discuss selected results in the "Overall Approach" section below.[3] More recently, scholars in accounting, finance, and law have focused their attention on the effects of institutional features (such as legal structure or judicial enforcement) on market frictions and their consequences for capital investment. For example, Bhattacharya et al. (2003) have shown that the cost of publicly traded equity across 34 countries is related to three different measures of earnings transparency. Francis et al. (2005) provide international evidence that dependence on external financing creates incentives for firms to undertake higher levels of voluntary accounting disclosure. However, the implication of these findings for firm-level investment is not straightforward. If equity financing were the only source of capital, then frictions in equity financing would probably lead to frictions in capital investing. However, firms can access multiple sources of financing. Thus, if one channel is inefficient (e.g., public equity financing), then other sources such as debt, private financing, state subsidies, and intra-group capital transfers, for example, can substitute. In equilibrium, there may not be any differences in investment efficiency at the firm level, but simply cross-sectional differences in financing patterns. Consistent with this view, Demirguc-Kunt and Maksimovic (2002) find no evidence that a country's relative reliance on bank versus stock market financing affects firms' access to external financing.

Other studies have considered relations between institutional features and capital investment at the industry level. For example, Wurgler (2000) measures cross-country capital allocation efficiency by industry-level elasticity of investment with respect to value added. He finds this measure to be positively related to the amount of firm-specific information available in domestic stock markets when measured by synchronicity, positively related to minority shareholder rights, and negatively related to state ownership. Rajan and Zingales (1998) show that industry growth is positively related to a financial development index.

Our paper advances the existing literature in several ways. First, instead of considering the effect of institutional features on industry growth (e.g., Rajan and Zingales 1998; Wurgler 2000) or on financial market development (e.g., LaPorta et al. 1997; Demirguc-Kunt and Maksimovic 1998), we focus on firm-level capital investment decisions. Second, we concentrate on accounting quality, a feature not examined in many prior studies (e.g., LaPorta et al. 1997; Demirguc-Kunt and Maksimovic 1998; Wurgler 2000). In addition, we consider the differences both between countries and between firms within a country. Third, our tests do not limit capital market frictions to their effects on capital rationing exclusively (e.g., Fazzari et al. 1988). Rather, we allow accounting quality to reduce either the lack or excess of cash. Finally, given the possibility that different sources of capital may substitute for each other, we do not limit our attention to any specific financing channel (e.g., Bhattacharya et al. 2003) and instead focus on capital investment behavior.

III. HYPOTHESIS DEVELOPMENT

Tobin (1969) theorizes that capital investment is a function of the ratio between the stock-market valuation of existing real capital assets and their current replacement cost. Yoshikawa (1980), Hayashi (1982), and Abel (1983) reconcile this theory with the neoclassical interpretation. In this framework, the marginal Q ratio, q, is the driver of capital investment policy. For example, Hayashi (1982) summarizes the model by stating, "once q is known ... the firm can decide the optimal rate of investment though the knowledge of the installation function ψ alone." He then proceeds to estimate this baseline model by regressing the investment rate on q. Fazzari et al. (1988) challenge this view and suggest that firms that are liquidity-constrained (i.e., cannot externally finance their investments) need to rely more on their internally generated funds.

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[3] See also Hubbard (1998) for a review and representative studies by Bagehot (1873), King and Levine (1993), Rousseau and Wachtel (1998), and Beck et al. (2000).

Hoshi et al. (1991) summarize two possible justifications for capital rationing. On the one hand, moral hazard models suggest that outside financing can dilute managements' ownership stakes, thereby exacerbating incentive problems that arise when managers control the firm but do not own it. This *ex post* incentive problem reduces the amount of capital supplied *ex ante*. On the other hand, Myers and Majluf (1984) propose an adverse selection model. They suggest that if managers are better informed than investors about a firm's prospects, they will try to sell overpriced securities. Rational investors will, in response, increase the cost of capital, thus decreasing the amount demanded. Therefore, in both cases, frictions operate to reduce the amount of external capital supplied to the firm. Firms that can generate cash internally are able to mitigate these effects, which causes capital investment to be correlated with the availability of internally generated funds.

There is presently a debate (see, for example, Fazzari et al. 2000; Kaplan and Zingales 1997, 2000) as to whether higher investment-cash flow sensitivity necessarily implies that firms are more financially constrained. In this paper, we remain agnostic on this issue since the problem may not be a lack, but rather an excess of cash. Capital investment can be correlated with internally generated funds simply because managers do not return to investors excess cash coming from rents (and quasi-rents) and other assets in place. Casual empiricism suggests the existence of such overinvestment by managers and several theoretical explanations have been proposed. For example, Jensen (1986) suggests that managers have incentives to grow their firms beyond their optimal size.[4] Whereas external financing subjects managers to monitoring and disciplining by capital providers, "financing projects internally avoids this monitoring and the possibility the funds will be unavailable," thereby allowing managers to overinvest. Blanchard et al. (1994) provide empirical support for this view. They consider what managers do when they receive a cash windfall that does not change the investment opportunity set (i.e., Tobin's Q). In perfect financial markets, managers should return the money to the capital suppliers. Contrary to this expectation, Blanchard et al. (1994) find that managers tend to invest in unrelated projects that typically fail.

Notice that all of these imperfections are predicated by the existence of information asymmetry between managers and outside suppliers of capital. If managers could commit to revealing all of their private information, then outsiders would not ration capital for fear of buying at an inflated price. Similarly, if higher quality accounting permitted perfect monitoring, then no agency problem would arise (see, for example, Antle and Eppen [1985] for a formal model of this idea). We would then be back to the baseline neoclassical model and internally generated cash flows would play no role in investment decisions. Our main hypothesis is predicated on this idea:

> **H1:** Higher quality accounting reduces the investment-cash flow sensitivity at the firm level.

We further predict, based on the reasoning above, that this effect should be stronger in economies where financing is largely provided through arm's-length transactions, for example, where the stock market is the dominant source of capital. On the other hand, in economies where creditors play a central role in financing, banks may be able to obtain information through alternate private channels (and thus mitigate adverse selection problems). They also may be in a better position to monitor the managers directly once the capital has been supplied (and thus mitigate the moral hazard). In this case, accounting quality should be less relevant to their decisions to supply capital and, as a result, we should observe a smaller effect of accounting quality on investment efficiency in credit-based economies. This leads to our second hypothesis:

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[4] For simplicity, we use "empire building" as our main way to motivate overinvestment. However, there are other models leading to a similar pattern such as overconfidence (e.g., Heaton 2002), the "quiet life" (Bertrand and Mullainathan 2003), and reputation (e.g., Baker 2000).

> **H2:** Higher quality accounting reduces the investment-cash flow sensitivity more in economies dominated by stock markets than in those dominated by creditors.

Notice that we do not form any prediction as to whether one type of economy is better than the other in mitigating frictions in the investment process.

IV. CASH FLOW SENSITIVITY: EMPIRICAL SETTING

Overall Approach

We employ both cross- and within-country tests, each with its own advantages. In the cross-country tests, average investment-cash flow sensitivities are estimated by country using firm-level data and then regressed on accounting quality and other variables of interest. This testing approach is likely to enhance the power of our tests by increasing both the cross-sectional variation and the magnitudes of financial frictions. It also enables us to consider the effects of institutional factors, such as creditor rights and legal origin, which cannot be easily studied in a single-country setting. As observed by Francis et al. (2005), "the United States is generally viewed as having a frictionless financial market with relatively easy access to external financing, and therefore it is not clear if U.S. findings necessarily generalize to countries with different financial or legal systems."

However, the cross-country approach also has some limitations. First, it is difficult to obtain sufficient time-series data from multiple countries to estimate firm-specific parameters. Second, different institutional settings and accounting rules make it difficult to compare firm-based measures across countries. For example, since asset revaluation is permitted in some countries but not in others, plant, property, and equipment and depreciation measures convey different meanings. To address these issues and to provide corroborating evidence that the hypothesized effect of accounting quality on investment efficiency operates *within* as well as across counties, we further examine two contrasting countries: the U.S. and Japan. Both countries have large economies with ample sources of external financing. However, in the U.S., stock markets play a central role in providing capital, while in Japan, *keiretsu* and banks are the major source of capital. As documented by Wurgler (2000), the ratio of stock market value to total credit market value is 0.64 for the U.S. but only 0.15 in Japan. Our focus on the U.S. and Japan allows us to assess whether the dominant form of financing influences the effect of accounting quality on investment decisions.

Samples and Data

To enhance comparability with prior research, we utilize data from the same set of 34 countries examined by Bhattacharya et al. (2003). For our cross-country tests, accounting and financial data are obtained from Compustat Global Vantage for the entire coverage period of the database (1993 to 2004). Accounting and financial data for the within-country tests are obtained for the U.S. from Compustat and for Japan from the PACAP database (1975 to 2001). The longer time series for the within-country samples allow us to estimate firm-specific parameters and utilize them in panel specifications. We focus on industrial firms and, as is customary, exclude utilities and firms in the financial, real estate, insurance, and public administration sectors.[5] Our resulting sample includes only publicly traded companies with access to public sources of capital. As these firms also can potentially access alternate sources of capital including bank loans, governmental financing, and private equity when public markets are inadequate, this provides a lower bound for the effects of accounting quality and other institutional features. Private firms with lesser access to capital and weaker outside monitoring would benefit relatively more from enhanced institutional features than would firms with more access to capital and stronger outside monitoring.

Proxies for Accounting Quality

For our cross-country tests, we use four country-specific measures of accounting quality. Three of these measures are adapted from Bhattacharya et al. (2003) and a fourth from Bushman et al. (2004). *Earnings aggressiveness* is based on the converse of conservatism as defined by Ball et al. (2000). Following Ball et al. (2000) and Bhattacharya et al. (2003), we expect countries with more aggressive accounting practices to exhibit lower firm-level capital investment efficiency than countries with less aggressive accounting practices. *Loss avoidance* is the ratio of the number of firms with small positive earnings minus the number of firms with small negative earnings divided by the sum of the two. This proxy is derived from Burgstahler and Dichev (1997) and Degeorge et al. (1999) who employ a similar measure for U.S. firms. Following Bhattacharya et al. (2003), we expect countries with greater loss avoidance to exhibit lower capital investment efficiency than countries with less loss avoidance. *Earnings smoothing* is based on the cross-sectional correlation between the change in accruals and the change in cash flows scaled by lagged total assets. Following Leuz et al. (2003) and Bhattacharya et al. (2003), we expect countries with a greater degree of earnings management to exhibit lower capital investment efficiency than countries with a smaller degree of earnings management. Following Bushman et al. (2004), *Timeliness* is based on the average ranking of "answers to the following interim reporting questions: frequency of reporting, count of disclosed items, and consolidation of interim reports." We expect countries with timelier reporting to exhibit higher capital investment efficiency than countries without timely reporting.

Because we are agnostic regarding which dimension of accounting quality plays a more significant role, we aggregate the four accounting quality measures into a summary index (AQ). In order to combine them meaningfully, we first form binary specifications. We do so by creating four binary variables for earnings aggressiveness, loss avoidance, earnings smoothing, and timeliness based on whether they have a better value than the median value in the cross-country sample. We then create *AQ* by summing up these four binary variables. Combining the four measures into an index has the advantage of reducing the effects of measurement errors in the individual accounting quality measures. This index also is more comparable with the index variables for creditor and shareholder rights developed by LaPorta et al. (1997) introduced as control variables below.[6]

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[5] We focus on industrial and commercial companies and, following prior studies, exclude U.S. firms with SIC codes between 4900–4999, 6000–6999, and above 9000, and those trading as American Depository Receipts (ADRs). We exclude Japanese firms with INDID codes starting with 05, 06, and 08.

[6] *AQ* takes values between 0 and 4. Below, we also control for financial disclosure as measured by an index created by the Center for International Financial Analysis and Research (CIFAR). The CIFAR index can potentially take values between 0 and 100. To make *AQ* and *CIFAR* more comparable, we also normalize *CIFAR* by forming a variable that takes a value between 0 and 4 based on quintiles of the distribution of *CIFAR*. Our results (untabulated) are not affected by this transformation.

In our within-country tests for the U.S. and Japan, we employ firm-specific accounting quality measures, AQ_{FS}, reflecting the uncertainty in accruals following Dechow and Dichev (2002). Francis et al. (2005) find that this measure of accounting quality is the most closely associated with their measure of the cost of capital. In Dechow and Dichev (2002), accruals quality is measured by the extent to which working capital accruals explain current-period, prior-period, and future-period operating cash flow realizations. The unexplained portion of the variation in working capital accruals is an inverse measure of accruals quality (with a larger unexplained portion implying lower quality accounting). Details of the estimation procedure are provided in the Appendix. We predict that higher quality accruals will be associated with higher capital investment efficiency especially in countries that rely on capital provided through arm's-length transactions.

Proxies for Investment-Cash Flow Sensitivity

We use two different but comparable procedures to measure investment-cash flow sensitivity in the cross-country and within-country tests. In the cross-country tests, we mirror the prior literature (e.g., Fazzari et al. 1988; Hoshi et al. 1991) by estimating investment-cash flow sensitivities after controlling for investment opportunities using Tobin's Q (proxied by the market-to-book equity ratio, *MTB*). Given the heterogeneity of our cross-country sample, outliers could induce nonlinear relations. A standard approach to alleviating this problem is to take the log of the variables, which we apply to market-to-book ratios. However, operating cash flows and investments often have negative values for which a log transformation is not defined. For them, we use an arctangent transformation that in effect "logs negative values."[7] The following model is estimated (using firm fixed effects) for each country:

$$I_{i,t}/K_{i,t-1} = \beta^{i}_{0} + \beta_1 OCF_{i,t}/K_{i,t-1} + \beta_2 MTB_{i,t} + \varepsilon_{i,t} \qquad (1)$$

where $I_{i,t}/K_{i,t-1}$ is capital investment scaled by the beginning-of-period capital for firm i in period t, $OCF_{i,t}$ is the operating cash flow, β_1 is our measure of investment-cash flow sensitivity, $\varepsilon_{i,t}$ is a normally distributed error term with a mean of zero, and it represents usual idiosyncratic errors.

This specification has two appealing features. First, it does not require a long time-series for each firm to estimate the parameters, which allows us to use pooled cross-sectional data to estimate an average investment-cash flow sensitivity for each country. This feature is important because Compustat Global Vantage contains a maximum of only 11 years of data. Second, this specification has been extensively used in the prior literature. However, it relies on *MTB* to proxy for Q. To the extent that *MTB* does not fully capture investment opportunities, cash flows may pick up measurement errors that are correlated across countries with accounting quality.

We address this concern in several ways. In our cross-country tests, we include future cash flows (Alti 2003) and we use an instrumental variables approach in some specifications. In our within-country tests, we use two databases, Compustat for the U.S. and PACAP for Japan, which provide longer time series. This allows us to use an alternate test that does not rely on estimating Q. By doing so, we have a natural robustness check for our cross-country estimations. We proxy for cash flow sensitivity of investment (*CFSI*) using the measure recently proposed by Hovakimian and Hovakimian (2005). $CFSI_0$ is measured as the difference between the cash-flow-weighted time-series average investment (*CFWAI*) of a firm and its un-weighted arithmetic time-series average investment (*AI*):

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[7] The arctangent function is approximately linear over the range (−1;1). In cases where the absolute value of investment or cash flows is less than net fixed assets, the data are changed little by the transformation. However, when the absolute values of these scaled variables increase beyond 1, the transformed value is bounded by $\pi/2$. This transformation provides a convenient way of winsorizing the data.

$$CFSI_{0,i,t} = CFWAI_{0,i,t} - AI_{i,t} = 1/n\Sigma_{s=1}{}^{t}\,[(CF_{i,s}/(\Sigma_{s=1}{}^{t}\,CF_{i,s})) * I_{i,s}] - 1/n\Sigma_{s=1}{}^{t}\,I_{i,s} \tag{2}$$

where *CF* is a firm's cash flow, *I* is the firm's investments (where both are scaled by beginning-of-period net capital), and *n* is the number of annual observations for firm *i*.[8]

The intuition behind this specification is that there should be no systematic difference between the weighted and un-weighted average investments for firms whose investment decisions are not affected by their available cash flows. On the other hand, the value of *CFSI* should be higher for firms that tend to invest more in years with relatively high cash flows and less in years with relatively low cash flows. Unlike the measure of cash flow sensitivity in Equation (1), we do not have to estimate a regression coefficient and, therefore, we can avoid the associated econometric estimation issues.

Estimation Procedures

Our cross-country tests regress the 34 investment-cash flow sensitivities from Equation (1) on our measure of accounting quality, other institutional features (disclosure quality, creditor rights, shareholder rights, judicial efficiency, and legal origin), and control variables such as the gross domestic product (*GDP*), the average tangibility of firm assets, and firm size, which are all standard in the literature (see the Appendix for variable definitions and estimation details). In the U.S. within-country tests, we also control for auditing quality and analyst coverage as well as other control variables suggested by Hovakimian and Hovakimian (2005) and described in the Appendix. In both the U.S. and Japan within-country tests, accounting quality, investment-cash flow sensitivity, and control variables are estimated using ten-year rolling windows. Standard errors are corrected for heteroscedasticity and for clustering of observations by firm.

V. CASH FLOW SENSITIVITY: EMPIRICAL RESULTS

Cash Flow Sensitivities and Between Countries Differences

Table 1 presents correlations among variables examined in the cross-country tests. Most correlations are reasonably small, suggesting that multicollinearity is not a significant concern. In particular, the correlation between *AQ* and *CIFAR* is only 0.13, suggesting that these two variables capture two related but differing constructs. However, shareholder rights (*SR*) is highly correlated with both *AQ* and *CIFAR,* as is legal origin (*LO*) with *AQ, CIFAR, SR*, and *CR*. We take these high correlations into consideration in our later tests.

Table 2 presents tests of H1 by regressing investment-cash flow sensitivities at the country level from Equation (1) on accounting quality (*AQ*) and other variables as described above. We report several specifications. In column 1, we enter only *AQ*. In columns 2 and 3, we control for other institutional features. Column 2 excludes *LO,* out of the concern that this variable is highly correlated with several other features. Finally, column 4 introduces controls for economy-wide factors. Results are consistent with our predictions. They

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[8] Variable definitions are detailed in the Appendix.

TABLE 1
Correlations among Summary Measures[a]

	AQ	*CIFAR*	*CR*	*SR*	*JE*
CIFAR	0.13 (0.48)				
CR	−0.17 (0.33)	0.04 (0.83)			
SR	0.31 (0.08)	0.39 (0.03)	0.22 (0.20)		
JE	−0.16 (0.37)	0.44 (0.01)	−0.15 (0.41)	−0.08 (0.66)	
LO	0.27 (0.12)	0.35 (0.04)	0.38 (0.03)	0.66 (0.00)	−0.11 (0.55)

[a] *AQ* is a measure of accounting transparency calculated as the sum of *Earnings aggressiveness, Loss avoidance, Earnings smoothing* (Bhattacharya et al. 2003), and *Timeliness* (Bushman et al. 2004). *CIFAR* is based on a disclosure index calculated by the Center for International Financial Analysis and Research. *CR* is a summary measure for creditor rights (as reported by La Porta et al. [1997]). *SR* is a summary measure for shareholder rights (as reported in La Porta et al. [1997]). *JE* is a measure of judicial efficiency based on the five measures found in La Porta et al. (1997). *LO* is an indicator variable that takes the value of 1 if a country's legal system has as a Common law origin, 0 otherwise (as reported by La Porta et al. [1997]).
The p-values are reported in parentheses.

indicate significant negative relations between *AQ* and the cash flow sensitivity of investment. When we turn our attention to the other institutional features (*CIFAR*, *JE*, *CR*, *SR*, and *LO*), we see that while most are associated with lower investment-cash flow sensitivity, only *CIFAR* and *CR* are consistently significant at conventional levels. This may be explained by a lack of power due to small sample size and it is not entirely unexpected given related findings in prior studies (e.g., DeFond and Hung 2004; Bushman et al. 2004).[9] Among the economic control variables, size has the strongest effect. Our results also appear to be economically significant, with the specification including all institutional factors (column 5) explaining 44 percent of the variation across countries in investment-cash flow sensitivity. *AQ* alone explains 14 percent of the variation (column 1).

As a robustness check, we also employ a "direct" estimation procedure that examines firm-year observations (untabulated). Specifically, we interact *AQ* with *OCF* in Equation (1) (we also include *AQ* in this specification). An advantage of this approach is that countries with larger economies comprise a larger share of the sample, so our sample is more representative of real economic activity. To prevent bias from measurement errors in Tobin's Q from driving our results, we estimate the model using a two-stage least squares (2SLS) procedure, whereby $MTB_{i,t}$ is instrumented by $MTB_{i,t-2}$ and $Kint_{i,t-3}$, lagged two and three periods, respectively, where *Kint* is the capital intensity defined as the ratio of net fixed assets to sales. Specification tests indicate that our instruments are both relevant and valid. We also apply ordinary least squares (OLS) techniques to ensure that our results do not hinge on a particular 2SLS specification. Standard errors in our pooled 2SLS specifications

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[9] DeFond and Hung (2004) report that international variations in shareholder rights do not affect CEO turnover. Bushman et al. (2004) find no relation between legal origin and their measure of financial transparency.

TABLE 2
Investment-Cash Flow Sensitivity
Cross-Country Sample[a]

Variable	**Predicted Sign**	**Dependent Variable: Investment-Cash Flow Sensitivity**			
		(1)	(2)	(3)	(4)
Intercept	+	0.527	1.725	1.558	1.994
		(5.07)	(3.66)	(3.12)	(3.69)
AQ	−	−0.120	−0.141	−0.129	−0.113
		(−2.74)	(−2.63)	(−2.40)	(−2.01)
CIFAR	−		−0.014	−0.013	−0.018
			(−2.29)	(−1.98)	(−2.52)
JE	−		−0.008	−0.010	−0.016
			(−0.63)	(−0.81)	(−1.19)
CR	−		−0.089	−0.071	−0.069
			(−2.48)	(−1.96)	(−1.88)
SR	−		0.038	0.056	0.075
			(0.92)	(1.24)	(1.52)
LO	−			−0.126	−0.207
				(−1.33)	(−1.86)
Log(GDP/capita)	−				−0.000
					(−1.37)
K-Intensity	−				0.002
					(0.06)
Size	−				−0.030
					(−3.39)
Number of observations		34	33	33	33
R^2		13.96%	42.05%	43.94%	55.32%

[a] *AQ* is a measure of accounting transparency calculated as the sum of *Earnings aggressiveness, Loss avoidance, Earnings smoothing* (Bhattacharya et al. 2003), and *Timeliness* (Bushman et al. 2004). *CIFAR* is based on a disclosure index calculated by the Center for International Financial Analysis and Research. *JE* is a measure of judicial efficiency based on the five measures found in La Porta et al. (1997). *CR* is a summary measure for creditor rights (as reported by La Porta et al. [1997]). *SR* is a summary measure for shareholder rights (as reported by La Porta et al. [1997]). *LO* is an indicator variable that takes the value of 1 if a country's legal system has as a Common law origin, 0 otherwise (as reported by La Porta et al. [1997]). *Log(GDP/Capita)* is gross domestic product per capita. *K-Intensity* is the average capital intensity for a given country defined as the ratio of net fixed assets to sales. *Size* is the average firm size in the country.

Regressions are estimated using ordinary least squares.

The t-statistics with standard errors adjusted for heteroscedasticity are reported in parentheses.

are adjusted for unspecified heteroscedasticity and serial-correlation using a two-step efficient generalized method of moments (GMM) estimator (Hayashi 2000).[10] Results are consistent with our results from the indirect specification. The z- (or t-) statistics range from −3.81 to −9.98 for the interaction between cash flows and our accounting quality index, and from −4.71 to −13.08 for the interaction with *CIFAR*. As expected, the z- (or t-) statistics are positive and range from 8.37 to 28.13 for cash flows. Finally, Alti (2003) suggests that there could be cash flow sensitivity even in an economy without frictions, because contemporaneous cash flows may better reflect information about short-term profitability than would the market-to-book equity ratio (*MTB*). To control for this possible effect, we add operating cash flow (*OCF*) at time $t+1$ and $t+2$ (scaled by fixed assets) or the difference in operating cash flows. Although this specification reduces the sample size (since we need two future years to estimate future profitability), results are qualitatively similar. Thus, our cross-country tests support H1 that higher-quality accounting reduces investment-cash flow sensitivity at the firm level.

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[10] Alternatively, we remove the year dummies and allow for clustering of observations by year to control for cross-correlation in the error term. Untabulated results are qualitatively similar.

Our second hypothesis predicts that higher quality accounting should reduce investment-cash flow sensitivity more in economies dominated by stock markets than in those dominated by creditors. To test this prediction, we partition our cross-country sample based on whether the ratio of stock market value to credit market value in the country is above or below the median in the sample. Consistent with our hypothesis, accounting quality is statistically significant in the subsample of equity-based economies (the t-statistic equals −2.06) but less than significant at conventional levels in the credit-based economies (the t-statistic equals −1.25).[11]

Cash Flow Sensitivities and Within-Country Differences: The U.S.

Results for the U.S. within-country tests reported in Table 3 are consistent with those for the international cross-sectional tests above. Column 1 reports the results for our panel specification where $CFSI_0$ is the dependent variable and AQ_{FS} is the only independent variable. In column 2, we introduce the set of control variables described above and in the Appendix. In column 3, following Hovakimian and Hovakimian (2005), we transform $CFSI_0$ into a binary variable that takes the value of 1 if $CFSI_0$ is greater than 0.05, and 0 otherwise, then use a logit specification with firm fixed effects to estimate the coefficients. All of the specifications consistently indicate that AQ_{FS} is associated with a reduction in the investment-cash flow sensitivity. This result is economically significant—an increase of AQ_{FS} by one standard deviation leads to a decrease of 30 percent in the average cash flow sensitivity.[12] Thus, higher accounting quality is associated with higher investment efficiency even in the most liquid market in the presence of predominant public equity financing. Our results (untabulated) also hold under alternate specifications, for example, when we use boot-strapped standard errors, add year dummies, use observations from the last year only, control for the analyst coverage, and measure $CFSI_0$ using cash flows lagged by one year. Auditing quality (*BigSix*) and analyst coverage (*AnalCov*) are significantly negative in (untabulated) univariate regressions, but these variables become insignificant in some multivariate specifications. The other control variables generally have the expected signs and are generally consistent with Hovakimian and Hovakimian (2005). For example, we would expect that larger, more profitable and more stable firms would be more transparent (see, among others, Lang and Lundholm 1993). Hence, these firms should exhibit less investment-cash flow sensitivity than smaller, less profitable and less stable firms. Consistent with these expectations, the coefficients on *Size*, *ROA*, *Z-score*, and *CFO* are negative, as expected, while the coefficients on $\sigma(CFO)$ and *M2B* are positive.

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[11] The magnitudes of the coefficients also are different (−0.15 for the equity-based sample and −0.09 for the credit-based sample). The small sample size prevents us from testing if these coefficient estimates differ significantly; pooling the two samples is also not feasible because of the high collinearity between the variable of interest in such a pooled sample. One should not necessarily conclude that accounting quality plays no role in credit-based economies for the same reason.

[12] We obtain this estimate by multiplying the value of the coefficient (1.62) by one standard deviation of AQ_{FS} (0.014) and then dividing by the average $CFSI_0$ (0.076).

TABLE 3
Investment-Cash Flow Sensitivity
U.S. Sample[a]

Variable	**Dependent Variable: Investment-Cash Flow Sensitivity**		
	(1)	(2)	(3)
AQ_{FS}	−2.48	−1.624	−27.57
	(−9.17)	(−6.24)	(−10.26)
BigSix		−0.111	−0.17
		(−2.41)	(0.43)
LogAsset		−0.028	−0.559
		(−6.85)	(−9.48)
Mkt-to-Book		0.048	1.257
		(5.92)	(11.76)
ROA		−0.434	−6.124
		(−3.92)	(−5.88)
$\sigma(CFO)$		0.639	11.916
		(5.18)	(12.14)
Z-score		−0.029	−0.401
		(−2.46)	(−2.91)
Tangibility		0.131	3.209
		(3.18)	(5.45)
R&D		0.000	−0.043
		(0.03)	(−0.22)
K-structure		−0.036	1.33
		(−1.09)	(3.59)
Mean K-structure		0.225	1.129
		(3.47)	(1.03)
CFOassets		−13.717	−48.145
		(−3.72)	(−1.60)
Slack		0.016	0.272
		(1.73)	(5.03)
Dividend		0.010	0.395
		(0.67)	(2.64)
Number of observations	28,353	28,353	12,420
R^2	68.24%	70.10%	

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[a] The dependent variable is a measure of cash flow sensitivity as defined by Hovakimian and Hovakimian (2005). AQ_{FS} is a measure of accounting quality as defined by Francis et al. (2005). *BigSix* is a dummy variable that takes the value of 1 if the firm is audited by a Big Six auditor (data # 159 between 10 and 89), 0 otherwise. *LogAsset* is the log of Compustat item 6. *Mkt-to-Book* is item 6 plus the product of item 25 and 199 minus item 60 and item 35, scaled by item 6. *ROA* is a measure of profitability calculated as the ratio of Compustat item 170 divided by item 6. $\sigma(CFO)$ is the standard deviation of cash flows from operations. *Z-score* is a measure of bankruptcy risk defined as 3.3 times item 170 + item 12 + one fourth of item 36 plus one half of the difference between item 4 and item 5, scaled by item 6. *Tangibility* is a measure of bankruptcy cost defined as the ratio of item 8 and item 6. *R&D* is a dummy variable that takes the value of 1 if item 46 is greater than 0, 0 otherwise. *K-structure* is item 9 scaled by data 9 plus the product of item 25 and 199. *Mean K-structure* is the mean of *K-structure* at the industry level. *CFOassets* is cash flows from operations divided by total assets. *Slack* is the ratio of item 1 and item 8. *Dividend* is a dummy variable that takes the value of 1 if item 21 or 127 is greater than 0, 0 otherwise.

Columns 1 and 2 are estimated using ordinary least squares. Column 3 is estimated using a logit specification. All specifications use firm fixed effects.

The t-statistics are reported in parentheses.

Cash Flow Sensitivities and Within-Country Differences: Japan

Untabulated results for the Japan within-country tests indicate that, contrary to the U.S. results, there is no statistically significant negative relation between accounting quality and investment-cash flow sensitivity. Although the Japanese sample size is smaller than the U.S. sample size (approximately 15,000 observations for Japan versus 30,000 for the U.S.), it is still sufficiently large to rule out statistical power as an explanation. This lack of statistical significance is consistent with H2 that accounting quality should have a smaller effect on investment efficiency in countries where bank financing and *keiretsu* are important sources of capital than in countries where equity is a dominant source of capital, because capital suppliers in credit-based economies have alternative avenues for reducing information asymmetry.[13]

VI. SUMMARY AND CONCLUSION

This study examines how accounting quality relates to firm-level capital investment efficiency. We consider several measures of accounting quality derived from prior research and two different methods for estimating investment-cash flow sensitivities. Our first hypothesis captures the intuition that higher quality accounting should enhance investment efficiency by reducing information asymmetry between managers and outside suppliers of capital. As argued in prior studies, information asymmetry can generate liquidity constraints (for example, as investors ration capital to protect themselves against adverse selection) or excess cash (for example, when it is difficult for owner-principals to monitor spending by manager-agents). Both situations generate inefficiencies in the investment process that accounting quality should mitigate. Our test results are consistent with this prediction both across and within countries under numerous alternative specifications. We find that the link between internally generated cash flows and investment is weaker when accounting quality is high. However, our tests do not address whether higher quality accounting operates primarily to mitigate the effect of a lack of cash or of its excess. This is a question we leave to future research.

Our second hypothesis is that the effect of higher quality accounting on investment-cash flow sensitivity should be stronger in economies where financing is largely provided through arm's length transactions, for example, where stock markets are dominant sources of capital. In contrast, in countries where creditors provide more capital, banks may be able to obtain information through alternative private channels and may be better positioned to monitor managers directly once capital is supplied. Thus, we predict a stronger (weaker) relation between accounting quality and capital investment efficiency in countries with predominant equity (bank) financing of firm-level capital investment. Our test results support this second hypothesis both across and within countries. Overall, our findings lend support to the argument that accounting quality is an institutional feature available to policy makers that enhances investment efficiency by mitigating investment-cash-flow sensitivity, and that this effect is stronger under conditions where capital is provided through arm's-length transactions.

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[13] The mean and the standard deviation of the winsorized investment-cash flow sensitivity estimates are 0.03 and 0.13, respectively, for Japanese firms versus 0.07 and 0.20, respectively, for U.S. firms. This suggests that homogeneity does not explain the Japanese results.

APPENDIX
Details of Variable Estimation

Accounting Quality

Cross-Country Sample

We create *AQ* by summing binary measures of four variables adapted from Bhattacharya et al. (2003) and Bushman et al. (2004) as described in the "Proxies for Accounting Quality" section above.

U.S. Sample

We first estimate the following model:

$$TCA^1_{j,t} = \varphi_{0,j} + \varphi_{1,j}CFO_{j,t-1} + \varphi_{2,j}CFO_{j,t} + \varphi_{3,j}CFO_{j,t+1} + u_{j,t}$$

where:

$TCA_{j,t} = \Delta CA_{j,t} - \Delta CL_{j,t} - \Delta Cash_{j,t} + \Delta STDEBT_{j,t}$ = total current accruals in year t;
$CFO_{j,t} = NIBE_{j,t} - TA_{j,t}$ = firm j's cash flow from operations in year t;
$NIBE_{j,t}$ = firm j's net income before extraordinary items (Compustat #18) in year t;
$TA_{j,t} = (\Delta CA_{j,t} - \Delta CL_{j,t} - \Delta Cash_{j,t} + \Delta STDEBT_{j,t} - DEPN_{j,t})$ = firm j's total accruals in year t;
$\Delta CA_{j,t}$ = firm j's change in current assets (Compustat #4) between year $t-1$ and year t;
$\Delta CL_{j,t}$ = firm j's change in current liabilities (Compustat #5) between year $t-1$ and year t;
$\Delta Cash_{j,t}$ = firm j's change in cash (Compustat #1) between year $t-1$ and year t;
$\Delta STDEBT_{j,t}$ = firm j's change in debt in current liabilities (Compustat #34) between year $t-1$ and year t; and
$DEPN_{j,t}$ = firm j's depreciation and amortization expense (Compustat #14) in year t.

We multiply the variance by minus one, so that a higher value of *AQ* corresponds to higher accounting quality.

In addition, we follow Francis et al. (2005) and use the model modified to include plant, property, and equipment (PPE) and change in revenues (scaled by average assets). McNichols (2002) proposes this extension, arguing that the changes in sales revenue and PPE are important in forming expectations about current accruals, over and above the effects of operating cash flows. She shows that adding these variables to the cross-sectional Dechow and Dichev (2002) regression significantly increases its explanatory power, thus reducing measurement error. However, the drawback of using this specification for our purpose is that it may include a mechanical link with our right-hand side variables by including capital investment in the regression. As a robustness check, we estimate a second model:

$$TCA^2_{j,t} = \psi_{0,j} + \psi_{1,j}CFO_{j,t-1} + \psi_{2,j}CFO_{j,t} + \psi_{3,j}CFO_{j,t+1} + \psi_{4,j}\Delta Rev_{j,t} + \psi_{5,j}\,PPE_{j,t} + \eta_{j,t}$$

where:

$\Delta Rev_{j,t}$ = firm j's change in revenues (Compustat #12) between year $t-1$ and year t; and
$PPE_{j,t}$ = firm j's gross value of PPE (Compustat #7) in year t.

To conserve space, we only tabulate the results from the first specification but results are very similar when we use the second one.

Japanese Sample

We follow a similar approach in our Japanese sample. Specifically, we use the following items from PACAP: INC9 and INC8 for *NIBE*; BAL6 for *CA*; BAL13 for *CL*; BAL1 for *Cash*; JAF34 for *STDEBT*; JAF74 for *DEPN*; INC1 for *Rev*; and the sum of JAF16, JAF17, JAF18, JAF19, JAF20, and JAF21 for *PPE*.

Variable I for Cash Flow Sensitivities

Cross-Country Sample

To calculate investment-cash flow sensitivities in the cross-country sample, we estimate the following model:

$$I_{i,t}/K_{i,t-1} = \beta^i{}_0 + \beta_1 OCF_{i,t}/K_{i,t-1} + \beta_2 MTB_{i,t} + \varepsilon_{i,t}$$

where $I_{i,t}$ is the investment in fixed assets for firm i in year t, defined as the change in net fixed assets (Global Vantage data76) plus depreciation expense (data11). $K_{i,t-1}$ is the total fixed assets for firm i in year $t-1$. $MTB_{i,t}$, our proxy for the Q ratio, is measured as the ratio of the sum of the market value of equity (data13 multiplied by data3) and the book value of debt divided (data106, data136, and data137 minus data138 and data139) by the book value of total asset (data89). One advantage of using the (quasi) market value of assets instead of their book value is that we do not drop from our sample firms with a negative equity book value. Consistent with prior literature (e.g., Hoshi et al. 1991), Q is calculated at the beginning of the period. Following previous studies,[14] $OCF_{i,t}$ is calculated as the sum of net income (data32) and depreciation expense (data11).

U.S. Sample

To calculate the cash flow sensitivity in our within-country samples, we estimate the model proposed by Hovakimian and Hovakimian (1985):

$$\begin{aligned} CFSI_{0,i,t} &= CFWAI_{0,i,t} - AI_{i,t} \\ &= 1/n\ \Sigma_{s=1}{}^{t}\ [(CF_{i,s}/(\Sigma_{s=1}{}^{t}\ CF_{i,s}) * I_{i,s}] - 1/n\ \Sigma_{s=1}{}^{t}\ I_{i,s} \end{aligned}$$

where n is the number of annual observations for firm i and t indicates the time period. I denotes investment, defined as capital expenditures (Compustat Item 128) divided by the beginning-of-period net capital. CF denotes cash flow and is defined as the sum of the income before extraordinary items (Compustat Item 18) and depreciation and amortization

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[14] See, among others, Morck et al. (1990), Hoshi et al. (1991), Whited (1992), Kaplan and Zingales (1997, 2000), and Hovakimian and Hovakimian (2005). We purposely do not correct our estimates of cash flows for any "working capital" effect for two reasons. First, we want to be consistent with prior literature on cash flow sensitivity of investment. Second, some firms lack the relevant information in our international database; thus, correcting our estimates of cash flows would introduce a selection bias. Our approach also has been used in prior international studies (e.g., Ball et al. 2000).

(Item 14), divided by the beginning-of-period net capital (Item 8). As an untabulated robustness check, we also consider a cash flow measure from Bushman et al. (2005) that excludes working capital accruals. Our results still hold. To avoid negative and extreme weighted values, we follow Hovakimian and Hovakimian (2005) and set the negative cash flows to zero. The variable is estimated using a ten-year rolling window. This specification does not require the estimation of Q. By doing so, we have a natural robustness check for the efficiency of our various statistical corrections in our cross-country sample.

In addition, since investment and cash flow are measured over an annual period, their exact timings may not coincide. In order to account for the possibility that investment may be financed with cash flows from the previous fiscal year, we also estimate *CFSI* based on *OCF*, which is lagged relative to investment:

$$\begin{aligned} CFSI_{-1,i,t} &= CFWAI_{-1,i,t} - AI_{i,t} \\ &= 1/n\ \Sigma_{s=1}{}^{t}\ [(CF_{i,s-1}/(\Sigma_{s=1}{}^{t}\ CF_{i,s-1})*I_{i,s}] - 1/n\ \Sigma_{s=1}{}^{t}\ I_{i,s} \end{aligned}$$

Results (untabulated) are not affected by using this alternative definition.

Japanese Sample

We follow a similar approach in our Japanese sample. Specifically, we use the following items from PACAP: change in *PPE* plus JAF74 scaled by beginning period *PPE* for *I* and *NIBE* plus JAF74, divided by the beginning-of-period net *PPE* for *CF*.

Control Variables in the Cash Flow Sensitivity Regressions

Cross-Country Sample

The institutional feature of primary interest is accounting quality, but we also consider financial disclosure, creditor rights (*CR*), shareholder rights (*SR*), judicial efficiency (*JE*), and legal origin (*LO*). As in many previous studies (e.g., Bhattacharya et al. 2003), we measure financial disclosure quality using an index created by the Center for International Financial Analysis and Research (CIFAR). *CR* and *SR* are based on the summary measures of creditor rights (*Cr5*) and shareholder rights (*Sr8*) obtained from La Porta et al. (1997). Our measure of judicial efficiency, *JE*, is the sum of the five proxies for judicial efficiency (*Rl1–Rl5*) found in La Porta et al. (1997). *LO* takes the value of 1 if the country has a legal origin rooted in the Common law tradition; 0 otherwise. We also control for economic characteristics: *Log(GDP/Capita)* is gross domestic product per capita as reported by the World Bank (2003); *K-Intensity* is the average capital intensity for a given country defined as the ratio of net fixed assets to sales; *Size* is the average firm size in the country measured as the log of the sum of the market value of equity and the book value of debt.

U.S. Sample

As proposed by Hovakimian and Hovakimian (2005), we control for size (log of Compustat item 6), market-to-book ratio (item 6 plus the product of items 25 and 199 minus item 60 and item 35, scaled by item 6), *ROA* (a measure of profitability calculated as the ratio of Compustat item 170 divided by item 6), dividend payout ratio (a dummy variable that takes the value of 1 if item 21 or 127 is greater than 0; 0 otherwise), leverage (item 9 scaled by data 9 plus the product of items 25 and 199), *Z-score* (a measure of bankruptcy risk defined as 3.3 times item 170 + item 12 + one fourth of item 36 plus one half of the difference between item 4 and item 5, scaled by item 6), the standard error of *CFO*, tangibility (a measure of bankruptcy cost defined as the ratio of item 8 and item 6), and *R&D* (a dummy variable that takes the value of 1 if item 46 is greater than 0; 0 otherwise).

In addition, we include two measures of financial slack, *CFOassets* (the ratio of *CFO* divided by item 6) and *Slack* (the ratio of item 1 and item 8). Finally, we control for auditing quality by including *BigSix*, a dummy variable taking the value of 1 if the firm is audited by a Big Six auditor (data # 159 between 10 and 89); 0 otherwise.

Generally, we expect bigger firms, with higher and more stable cash flows, more tangible assets, and fewer investment opportunities to be less sensitive to internally generated cash flows. We also control for analyst coverage in an untabulated robustness check. To measure analyst coverage, we use data obtained from the I/B/E/S Historical Summary File. We use *AnalCov*, the log of the number of analysts (plus 1, since the log of 0 is not defined) reported by I/B/E/S as covering the firm, as our proxy for analyst coverage. For each year, we set the number of analysts following a firm as the maximum number of analysts who make annual earnings forecasts in any month over a 12-month period. We assume that firms not covered by I/B/E/S are not covered by analysts.

Japanese Sample

We follow a similar approach in our Japanese sample. Specifically, we use the following items from PACAP: log of BAL9 for *size*; the sum of BAL9 and market value minus BAL21 scaled by BAL9 for *market-to-book* ratio; the sum of INC9, INC8, and INC7 scaled by BAL9 for *ROA*; MKT1 for calculating *dividend payout ratio*; the ratio of BAL14 divided by market value and BAL14 for *leverage*; 3.3 times the sum of INC9, INC8, and INC7 plus INC1 plus 0.25 time BAL20 plus 0.5 times the difference between BAL6 and BAL13, scaled by BAL9 for *Z-score*; the ratio *CFO* over *INC1* for *tangibility*; the ratio of *CFO* divided by *BAL9* for *CFOassets*; and the ratio of BAL1 scaled by the sum of JAF16 through 21 and 23 for *Slack*.

REFERENCES

Abel, A. B. 1983. Optimal investment under uncertainty. *American Economic Review* 73 (1): 228–233.

Alti, A. 2003. How sensitive is investment to cash flow when financing is frictionless? *Journal of Finance* 63 (2): 707–722.

Antle, R., and G. Eppen. 1985. Capital rationing, and organizational slack in capital budgeting. *Management Science* 31 (2): 163–174.

Bagehot, W. 1873. *Lombard Street: A Description of the Money Market.* Reprint 1999. Hoboken, NJ: Wiley.

Baker, M. 2000. Career concerns, and staged investment. Working paper, Harvard University.

Ball, R., S. P. Kothari, and A. Robin. 2000. The effect of international institutional factors on properties of accounting earnings. *Journal of Accounting and Economics* 29: 1–51.

Beck, T., R. Levine, and N. Loayza. 2000. Finance and the source of growth. *Journal of Monetary Economics* 46: 31–77.

Bertrand, M., and S. Mullainathan. 2003. Enjoying the quiet life? Corporate governance and managerial preferences. *Journal of Political Economy* 111 (5): 1043–1075.

Bhattacharya, U., H. Daouk, and M. Welker. 2003. The world price of earnings opacity. *The Accounting Review* 78 (3): 641–678.

Blanchard, O., F. Lopez-de-Silanez, and A. Shleifer. 1994. What do firms do with cash windfalls? *Journal of Financial Economics* 36 (3): 337–360.

Burgstahler, D., and I. Dichev. 1997. Earnings management to avoid earnings decreases and losses. *Journal of Accounting and Economics* (24): 99–126.

Bushman, R., J. Piotroski, and A. Smith. 2004. What determines corporate transparency? *Journal of Accounting Research* 42 (2): 207–252.

———, A, Smith, and F. Zhang. 2005. Investment-cash flow sensitivities are really capital investment-working capital investment sensitivities. Working paper, University of North Carolina.

Claessens, S., S. Djankov, and L. H. P. Lang. 2000. The separation of ownership and control in East Asian corporations. *Journal of Financial Economics* 58 (1): 81–112.

Dechow, P., and I. Dichev. 2002. The quality of accruals and earnings: The role of accrual estimation errors. *The Accounting Review* 77 (Supplement): 35–59.

DeFond, M., and M. Hung. 2004. Investor protection and corporate governance: Evidence from worldwide CEO turnover. *Journal of Accounting Research* 42 (2): 269–312.

Degeorge, F., J. Patel, and R. Zeckhauser. 1999. Earnings management to exceed thresholds. *Journal of Business* 72 (1): 1–33.

Demirguc-Kunt, A., and V. Maksimovic. 1998. Law, finance, and firm growth. *Journal of Finance* 53: 2107–2137.

———, and ———. 2002. Funding growth in bank-based and market-based financial systems: Evidence from firm-level data. *Journal of Financial Economics* 65 (3): 337–363.

Fazzari, S., R. G. Hubbard, and B. Petersen. 1988. Investment and finance reconsidered. *Brookings Papers on Economics Activity*: 141–195.

———, ———, and ———. 2000. Investment-cash flow sensitivities are not useful: A comment on Kaplan and Zingales. *The Quarterly Journal of Economics* 115 (2): 695–705.

Francis, J., I. Khurana, and R. Pereira. 2005. Disclosure incentives and effects on cost of capital around the world. *The Accounting Review* 80 (3): 1125–1162.

Gorton, G., and F. Schmid. 2000. Universal banking and the performance of German firms. *Journal of Financial Economics* 20 (1): 29–80.

Hayashi, F. 1982. Tobin's marginal *q* and average *q*: A neoclassical interpretation. *Econometrica* 50 (1): 213–224.

———. 2000. *Econometrics*. Princeton, NJ: Princeton University Press.

Healy, P., and K. Palepu. 2001. Information asymmetry, corporate disclosure, and the capital markets: A review of the empirical disclosure literature. *Journal of Accounting and Economics* 31 (1-3): 405–440.

Heaton, J. 2002. Managerial optimism and corporate finance. *Financial Management* 31 (2): 33–45.

Hoshi, T., A. Kashyap, and D. Scharfstein. 1991. Corporate structure, liquidity and investment: Evidence from Japanese industrial groups. *The Quarterly Journal of Economics* 106 (1): 33–60.

Hovakimian, A., and G. Hovakimian. 2005. Cash flow sensitivity of investment. Working paper, City University of New York.

Hubbard, R. G. 1998. Capital-market imperfections and investment. *Journal of Economic Literature* 36 (1): 193–225.

Jensen, M. 1986. Agency costs of free cash flow, corporate finance, and takeovers. *American Economic Review* 76 (2): 323–329.

Kaplan, S., and L. Zingales. 1997. Do financing constraints explain why investment is correlated with cash-flows. *The Quarterly Journal of Economics* 112 (2): 169–215.

———, and ———. 2000. Investment-cash flow sensitivities are not valid measures of financing constraints. *The Quarterly Journal of Economics* 112 (2): 707–712.

King, R., and R. Levine. 1993. Finance and growth: Schumpeter might be right. *The Quarterly Journal of Economics* 108 (3): 717–737.

Lang, M., and R. Lundholm. 1993. Cross-sectional determinants of analysts' ratings of corporate disclosures. *Journal of Accounting Research* 31 (2): 246–271.

La Porta, R., F. Lopez-De-Silanes, A. Shleifer, and R. Vishny. 1997. Legal determinants of external finance. *Journal of Finance* 52 (3): 1131–1150.

Leuz, C., D. Nanda, and P. Wysocki. 2003. Earnings management and institutional factors: An international comparison. *Journal of Financial Economics* 69 (3): 101–122.

Morck, R., A. Shleifer, and R. Vishny. 1990. The stock market and investment: Is the market a sideshow? *Brookings Papers on Economic Activity* 1: 157–215.

Myers, S. 1984. The capital structure puzzle. *Journal of Finance* 39 (3): 575–592.

———, and N. Majluf. 1984. Corporate financing and investment decisions when firms have information that investors do not have. *Journal of Financial Economics* 13 (2): 187–221.

Rajan, R., and L. Zingales. 1998. Financial dependence and growth. *American Economic Review* 88: 559–586.

———, and ———. 2000. The great reversals: The politics of financial development in the 20th Century. Working paper, OECD.

Rousseau, P., and P. Wachtel. 1998. Financial intermediation and economic performance: Historical evidence from five industrialized countries. *Journal of Money, Credit and Banking* 30 (4): 657–678.

Tobin, J. 1969. A general equilibrium approach monetary theory. *Journal of Money, Credit and Banking* 1: 15–29.

Yoshikawa, H. 1980. On the *"q"* theory of investment. *American Economic Review* 70 (4): 739–743.

Whited, T. 1992. Debt, liquidity constraints and corporate investment: Evidence from panel data. *Journal of Finance* 47: 1425–1460.

World Bank. 2003. *World Bank Indicators*. Washington, D.C.: World Bank.

Wurgler, J. 2000. Financial markets and the allocation of capital. *Journal of Financial Economics* 58: 187–214.

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The Importance of Reporting Incentives: Earnings Management in European Private and Public Firms

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ABSTRACT: This paper examines how capital market pressures and institutional factors shape firms' incentives to report earnings that reflect economic performance. To isolate the effects of reporting incentives, we exploit the fact that, within the European Union, privately held corporations face the same accounting standards as publicly traded companies because accounting regulation is based on legal form. We focus on the level of earnings management as one dimension of accounting quality that is particularly responsive to firms' reporting incentives. We document that private firms exhibit higher levels of earnings management and that strong legal systems are associated with less earnings management in private and public firms. We also provide evidence that private and public firms respond differentially to institutional factors, such as book-tax alignment, outside investor protection, and capital market structure. Moreover, legal institutions and capital market forces often appear to reinforce each other.

Keywords: *international accounting; earnings management; private companies; legal system; accounting harmonization; earnings properties.*

JEL Classification: *G14, G15, G30, G32, K22, M41.*

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I. INTRODUCTION

Much of the international accounting debate has focused on accounting standards per se, which are often viewed as the primary input for high-quality accounting (e.g., Levitt 1998). Consistent with this view, harmonization efforts within the European Union have largely focused on eliminating differences in accounting standards across countries or adopting a common set of standards (e.g., Van Hulle 2004). However, it is important to also examine the role of institutional factors and capital market forces in shaping firms' incentives to report informative earnings. The application of standards involves judgment and underlying measurements are often based on private information. The resulting discretion can be used by corporate insiders either to make reported earnings more informative about the firm's economic performance or to serve other and less benign interests. For this reason, reporting incentives and the forces shaping them are likely to play an integral role for accounting quality. While this insight is not new (e.g., Watts and Zimmerman 1986), it is often overlooked in international standard setting.[1]

To empirically document the importance of reporting incentives, we examine the properties of reported earnings of private and public firms in the European Union (EU). The European setting provides a unique opportunity because EU accounting regulation is based on a firm's legal form, rather than listing status. Thus, private limited companies face largely the same accounting standards as publicly traded corporations, but are subject to very different capital market forces. This feature allows us to study the role of reporting incentives and the demand for information created by public equity markets, both of which are issues of fundamental economic importance. Several prior studies suggest that equity markets have a negative impact on accounting quality (e.g., Teoh et al. 1998a, 1998b; Beatty et al. 2002). However, these studies are either limited to specific industries or major corporate events. Thus, the first-order effect of public equity markets on accounting quality is still an open issue and one that we can analyze in the EU setting. The setting also allows us to examine the effects of cross-country variation in firms' institutional environments and to explore how legal institutions and capital market forces interact in shaping the reporting behavior of private and public firms.

We hypothesize that capital markets as well as critical aspects of a firm's institutional environment determine the role of earnings, e.g., its importance in resolving information asymmetries and in communicating with outside parties (e.g., Watts and Zimmerman 1986; Ball 2001). This role in turn influences how corporate insiders use reporting discretion, which crucially determines the properties of reported earnings.

The use of discretion and the resulting informativeness (or opacity) of earnings are difficult to measure because true economic performance is unobservable. Previous work in international accounting has often focused on the conservatism dimension of accounting quality, i.e., the extent to which losses are incorporated into earnings on a timely basis (e.g., Ball et al. 2000; Ball et al. 2003; Ball et al. 2005; Bushman and Piotroski 2006; Peek et al. 2006). As accounting quality is a broad concept with multiple dimensions, it is important to extend empirical results beyond the conservatism dimension. In this study, we focus on another dimension of accounting quality, namely the degree of earnings management.[2] We rely on an earnings management index suggested by Leuz et al. (2003), which is based on four different proxies. The underlying measures are designed to capture a variety of earnings management practices, such as earnings smoothing and accrual manipulations.[3] We also conduct sensitivity analyses using alternative earnings management metrics similar to those used by Lang et al. (2003) and Lang et al. (2006) as well as measures of conservatism. Our results are similar across these measures and consistent with the idea that all these measures capture aspects of earnings informativeness.

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[1] More frequently, standard setters recognize the importance of enforcement institutions as a determinant of earnings quality. For instance, Tweedie and Seidenstein (2005) note that a "sound financial reporting infrastructure" must consist of "an enforcement or oversight mechanism that ensures that the principles as laid out by the accounting and auditing standards are followed."

[2] There are other dimensions of accounting quality, e.g., how well earnings predict future cash flows. See Dechow and Schrand (2004) for a definition and a synthesis of the earnings quality literature.

We document substantial variation in earnings informativeness across private and public firms from 13 EU countries, despite decades of accounting harmonization. More importantly, we show that differences in firms' reporting incentives explain this variation. In particular, we find that earnings management is more pervasive in private firms than in publicly traded firms. Thus, contrary to recent allegations that capital markets exacerbate incentives to manage earnings, our findings suggest that the first-order effect of public equity markets is to improve earnings informativeness, either by providing incentives to make earnings more informative or, alternatively, by screening out firms with less informative earnings in the going public process. We also document that earnings management is more pronounced in countries with weaker legal systems and enforcement. This finding holds for both private and public firms and highlights the central importance of enforcement mechanisms.

Moving beyond general characterizations of legal systems, we explore the interaction between market forces and other institutional variables that have the potential to differentially affect private and public firms: (1) the degree of alignment between financial and tax accounting, (2) remaining differences in EU accounting rules, (3) the level of required disclosures in public securities offerings and associated enforcement, (4) the level of minority-shareholder protection, and (5) the structure and activity of capital markets.

The analysis of these interactions supports our earlier finding that capital market forces, by and large, improve the informativeness of earnings. We document that stronger tax alignment is associated with more earnings management, but this effect is mitigated for public firms. We find that legal institutions designed to facilitate equity financing at arm's length in public markets, such as strong minority-shareholder rights and extensive disclosure requirements, are associated with lower levels of earnings management primarily for publicly traded firms, suggesting that markets and institutions reinforce each other. Finally, in countries with large and highly developed equity markets, public firms engage in even less earnings management, which again suggests that strong capital markets and arm's-length financing improve earnings informativeness.

Our study contributes to the literature in several ways. First, by comparing the reporting behavior of private and public firms, we shed light on the first-order effects of public equity markets. Prior evidence on differences in earnings quality between public and private firms is conflicting, and either confined to a particular country (Vander Bauwhede et al. 2003; Ball and Shivakumar 2005), a single regulated industry (Beatty and Harris 1999; Beatty et al. 2002) or specific corporate events (Teoh et al. 1998a, 1998b). Our paper adds to this limited body of work by examining a large sample of public and private firms across many industries and countries, and outside of specific corporate events. Our findings generally support the notion that the first-order effect of capital market forces is to improve reporting quality.

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[3] We use earnings management as a measure of accounting quality for several reasons. First, earnings management measures should be particularly responsive to the use of discretion and firms' reporting incentives, increasing the power of our tests. Second, the measures are widely used and have been shown to produce plausible rankings of earnings informativeness (e.g., Lang et al. 2003; Wysocki 2004; Lang et al. 2006). Third, by showing the robustness of our findings across multiple measures, we are able to minimize concerns that results are attributable to spurious relations of any particular measure.

Second, we contribute to a fairly recent literature that analyzes the effects of capital market incentives on the properties of reported earnings but is limited to public firms (e.g., Ball et al. 2005; Bushman and Piotroski 2006). Data on private and public firms allows us to shed light on the interplay of market forces stemming from public equity with legal institutions that facilitate arm's-length financing in public markets, such as minority-shareholder protection and securities regulation. Moreover, we can study the extent to which equity-based financial systems and highly developed capital markets explain differences in reporting behavior. Our findings suggest that these institutions and capital market characteristics tend to reinforce the first-order effect of publicly traded equity.

Third, there is little evidence on how institutional factors shape the reporting behavior of private firms. With the exception of a concurrent study by Peek et al. (2006) that looks at earnings conservatism across a sample of private and public EU firms, prior work documents institutional effects on firms' reporting behavior using public firms only (e.g., Ali and Hwang 2000; Ball et al. 2000; Fan and Wong 2002; Ball et al. 2003; Leuz et al. 2003; Bushman et al. 2004). While it is safe to assume that institutions affect private firms as well, it is *a priori* not obvious in what way. We show that there are some institutions like legal enforcement that matter to both public and private firms. But there are others, such as tax alignment, where the negative effect on reporting quality differs across public and private firms, displaying the mitigating influence of market forces.

Finally, our paper contributes to the regulatory issue of accounting harmonization (e.g., Gernon and Wallace 1995; Saudagaran and Meek 1997) and the debate on accounting convergence (Joos and Lang 1994; Land and Lang 2002; Bradshaw and Miller 2005; Joos and Wysocki 2004). Our paper provides evidence supporting the conjecture that effective accounting harmonization is unlikely to be achieved by accounting standards alone (e.g., Ball 2001).

The paper is organized as follows. Section II develops our hypotheses. Section III describes the data and the research design. In Section IV, we present the evidence on the relation between listing status, legal enforcement, and earnings management. In Section V, we analyze the role of additional institutional factors that have the potential to differentially affect private and public firms. Section VI concludes.

II. THE EFFECT OF CAPITAL MARKETS AND INSTITUTIONAL FACTORS ON REPORTED EARNINGS

Our analysis is based on the recognition that accounting standards provide considerable discretion to firms in preparing their financial statements. Corporate insiders can use their private information to report earnings that more accurately reflect firm performance and are more informative to outside parties. However, if earnings play a minor role in communicating performance to outsiders, then insiders are unlikely to do so. Instead, reporting choices may be governed by other considerations, e.g., by the desire to minimize taxes or determine dividend payments. Moreover, corporate insiders can use reporting discretion to hide poor economic performance, achieve certain earnings targets, or avoid covenant violations. Given insiders' information advantage, it is difficult to constrain such behavior.[4]

These arguments suggest that factors that define the role of earnings and shape firms' reporting incentives play an integral role in determining the properties of reported earnings, such as their informativeness to corporate outsiders (e.g., Ball et al. 2000).[5] We argue that capital market forces and the home country's institutional features are such factors.

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[4] Note, however, that this is not just an enforcement issue. Even when holding the level of enforcement constant, reporting behavior will differ as long as firms have some discretion and their incentives differ.

Capital Market Forces

Privately held firms and those with publicly traded equity securities face very different demands for accounting information. External financing in public equity markets creates the demand for information that is useful in evaluating and monitoring the firm. Arm's-length equity investors do not have private access to corporate information and rely heavily on *public* information, such as financial statements and reported earnings. If the quality of this information is poor, then outside investors will be reluctant to supply capital to firms. As a result, publicly traded firms have stronger incentives to provide financial statements that help outsiders assess economic performance. In addition, the going public process may screen out firms with less informative earnings that are difficult to evaluate for outside investors. Thus, regardless of the mechanism, being public is likely to be associated with higher reporting quality.

In contrast, privately held firms have relatively concentrated ownership structures and hence can efficiently communicate among shareholders via *private* channels. Because financial statements and reported earnings assume a less important role in communicating firm performance, private firms have relatively fewer incentives to report informative earnings.[6] Accordingly, private firms can place greater relative weight on different roles for reported earnings than can public firms. For instance, it is of a lesser concern to private firms that managing earnings to minimize taxes may make earnings less informative to outsiders. Alternatively, earnings can be used in determining dividends and other payouts to firms' stakeholders. As in Ball and Shivakumar (2005), we argue that these other uses are likely to render earnings of private firms less informative.

While these arguments suggest that the first-order effect of public securities markets is to create incentives to report earnings that reflect economic performance, we recognize that there are trade-offs and potentially important countervailing effects. For instance, controlling insiders in public firms might expropriate outside investors by consuming large private control benefits. As an attempt to hide these activities and prevent outsider intervention, they could mask firm performance by managing reported earnings (Leuz et al. 2003).[7] However, analogous arguments can be made for private firms that are highly levered and in financial distress. The fear of creditor interference in this situation and the subsequent loss of private control benefits can also create incentives to obfuscate true performance.[8] Moreover, insiders in private firms may attempt to transfer assets out of the firm, effectively expropriating the creditors, which in turn creates similar hiding incentives as described for public firms. Thus, there are specific situations in which both private and public firms have strong incentives to misrepresent economic performance, but it is not clear whether there is a differential effect and how commonly these situations arise.

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[5] This logic has also been exploited in the accounting choice and earnings management literature. See Watts and Zimmerman (1986), Healy and Wahlen (1999), and Dechow and Skinner (2000).

[6] In addition, private firms may have incentives to obfuscate firm performance because the EU requires them to file financial statements to the corporate registrar. In particular, family-owned private firms may try to hide true firm performance from employees and the general public.

[7] Other examples of forces that could give rise to more earnings management in public firms are managerial compensation contracts, debt covenants, particularly in public debt agreements, or political pressure. See Healy and Wahlen (1999) for a survey of the empirical evidence on these forces.

[8] Extensive creditor rights in many European countries make the threat of creditor intervention even more real for private firms. While it is common in some European countries that bank representatives sit on supervisory boards of public firms, making them effectively corporate insiders, banks rarely assume this role in private firms, for the most part because supervisory boards are less common. Therefore, reported performance is likely to be an important trigger for lender intervention.

Another reason why capital markets can create incentives to manage reported earnings is related to earnings targets set for public firms. Beatty et al. (2002) argue that small investors in stock markets are more likely to rely on simple heuristics than fairly sophisticated private investors (e.g., banks), which makes public firms more likely to engage in accounting manipulations to meet or beat earnings targets. In a similar vein, Fischer and Stocken (2004) argue that analyst forecasts can create incentives to manipulate earnings. Consistent with these claims, Beatty and Harris (1999) and Beatty et al. (2002) present evidence from the U.S. banking industry suggesting that public banks engage in more earnings management than their private counterparts. However, as the banking industry is heavily regulated, which may result in unusual reporting incentives, it is not clear that the results extend to a more general setting. The same holds for earnings management documented around firms' seasoned equity offerings or initial public offerings (Teoh et al. 1998a, 1998b).

Thus, it is ultimately an empirical question to what extent these opposing effects prevail and whether capital markets push public firms to report earnings that are more informative about firm performance than do private firms. Accordingly, we formulate our tests as two-sided.

Enforcement Quality and Other Institutional Features

The home country's institutional framework also shapes reporting incentives. Prior work by La Porta et al. (1998) and La Porta et al. (1999) show that certain institutional structures encourage external equity financing and dispersed ownership structures, which in turn determines the role of earnings in reducing information asymmetries. Prior studies focus on summary measures for these institutional structures (e.g., legal origin) or on institutions that are primarily geared toward publicly traded firms (e.g., outside investor protection) and demonstrate that institutional differences influence the reporting behavior of public firms (e.g., Ball et al. 2000; Leuz et al. 2003; Bushman et al. 2004).

There is little evidence on how private firms respond to institutional factors. To address this void, we examine a number of key institutional features. The first factor is the quality of legal enforcement. Without proper enforcement, legal rules remain largely ineffective. Thus, we hypothesize that firms in countries with weak legal enforcement are more likely to abuse discretion afforded by the accounting rules. This logic is fairly general and should apply to both private and public firms. Holding everything else constant, a firm that operates in a country where courts are inefficient and the chance of prosecution is remote seems more likely to engage in earnings manipulations than in a country where courts are efficient.

We also examine four factors where we might expect a differential effect on public and private firms: (1) financial accounting and tax alignment, (2) differences in accrual accounting rules, (3) securities regulation and minority-shareholder protection, and (4) capital market structure. These are discussed in more detail below.

First, we analyze the effect of tax accounting on reported earnings numbers. Ball et al. (2000) hypothesize that the link between financial and tax accounting can play an important role in firms' reporting behavior.[9] Specifically, a close link between reported earnings and taxable income is expected to muddle firms' incentives to report true economic performance (e.g., Alford et al. 1993). Moreover, it is plausible that the degree of tax alignment of financial accounting has a differential effect on private and public firms. As private firms are less reliant on earnings to communicate firm performance, it is less of a concern to private firms if they make earnings less informative in the process of minimizing taxes. In contrast, we expect public firms to be relatively less likely to sacrifice earnings informativeness.

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[9] However, there is little direct evidence supporting this hypothesis. The recent work by Coppens and Peek (2005) is a notable exception, suggesting that taxes could play a role in observed earnings distributions as well as firms' tendency to avoid losses.

Second, we examine the effect of accounting rules that are designed to produce timely and informative reported earnings. In principle, such accrual rules should have a positive effect on earnings informativeness of all firms, provided that they are properly enforced. However, accounting rules that make more extensive use of accruals provide more reporting discretion to firms and are more difficult to enforce. As a consequence, we hypothesize that the effect of accrual rules depends on firms' reporting incentives and, therefore, is likely to differ across private and public firms. That is, we expect accounting rules that make heavier use of accruals to be associated with less earnings management for public firms than private firms.

Third, we examine whether stricter disclosure rules in securities offerings and associated enforcement make it harder for firms to engage in earnings management and create incentives to reveal economic performance. However, as these rules apply only to firms with publicly traded securities, we expect to observe little or no effect for privately held firms. Similarly, strong minority-shareholder protection rules are designed to facilitate equity financing at arm's length in public markets and hence are expected to reduce earnings management primarily for publicly traded firms.

Fourth, we consider the effects of two important aspects of capital market structure, the extent to which the financial system is relatively more equity-market- or bank-based and the degree of financial development. Equity-based financial systems rely more on arm's-length financing and hence create demand for high-quality financial statements. In contrast, bank-based financial systems support relationship lending, which leads to more financial information being privately communicated, rather than via financial statements (see also, Bushman and Piotroski 2006). More generally, market forces faced by public firms are likely to be stronger when a country's capital markets are active and highly developed. Thus, we expect that these two capital market features are negatively associated with earnings management primarily for the public firms.

III. RESEARCH DESIGN AND DATA

Accounting Harmonization in the European Union

The European setting provides a unique opportunity to analyze the role of reporting incentives. First, there is a substantial range of institutional differences across Europe. For example, the U.K. is generally viewed as an outsider economy with legal institutions and accounting regulations designed to meet the demands of arm's-length financing in debt and equity markets (see e.g., Ball et al. 2000; Leuz et al. 2003). On the other end of the spectrum, Germany and Italy are typically referred to as insider economies where firms establish close relationships with banks and other financial intermediaries, and legal institutions and accounting rules are designed to facilitate debt contracting and relationship financing. The Netherlands and the Scandinavian countries are generally viewed as being somewhere in the middle.

Second, accounting regulation within the European Union is not based on listing status. Privately held companies with limited liability face largely the same accounting standards as publicly traded corporations. Thus, within a given country, accounting standards are held constant across the two sets of firms. Moreover, accounting standards have been formally harmonized across EU countries for many years (Van Hulle 2004).

The cornerstones of EU accounting harmonization are the Fourth and the Seventh Directives. The Fourth Directive applies to some five million limited liability companies in the EU, whether they are publicly traded or privately held. It requires these firms to prepare audited financial accounts according to the principles laid out in the directive and to provide publicly accessible financial statements to the corporate registers. The Seventh Directive complements these harmonization efforts with requirements on when and how firms must prepare consolidated financial accounts. Thus, the two directives effectively prescribe a *common* set of accounting rules for private *and* public corporations with respect to their consolidated and unconsolidated (or parent-only) financial statements.[10]

Thus, the European setting is unique insofar as it provides substantial within-country and cross-country variation in capital market forces and institutional incentives while holding the accounting rules largely constant. Our study exploits this variation by explicitly linking it to the properties of reported earnings.

Proxies for Earnings Management

Our hypotheses call for measures that capture the extent to which firms use reporting discretion to make earnings more (or less) informative about the underlying economic performance. However, it is impossible to directly observe how firms use discretion to influence the informativeness of earnings. We therefore focus on the level of earnings management as a proxy for the extent to which firms use reporting discretion to reduce the informativeness of earnings. Note that we do not claim that firms always use discretion in a way to obfuscate economic performance.[11] However, we do rely on the assumption that more extreme realizations of our earnings management measures for a large set of private (or public) firms in an industry and country over several years suggest that reported earnings are less likely to reflect economic performance.

Following Leuz et al. (2003) and drawing on prior accounting research (e.g., Healy and Wahlen 1999; Dechow and Skinner 2000), we compute four different proxies capturing a range of earnings management activities: (1) the tendency of firms to avoid small losses, (2) the magnitude of total accruals, (3) the smoothness of earnings relative to cash flows and (4) the correlation of accounting accruals and operating cash flows. We recognize that these proxies are not perfect and indicate earnings management only in a relative sense, but relative measures meet the demands of our design. Moreover, recent studies using these proxies suggest that they yield country rankings that are consistent with widespread perceptions of earnings informativeness, and that they behave in a plausible fashion (e.g., Lang et al. 2003; Lang et al. 2006; Wysocki 2004).

Burgstahler and Dichev (1997) and Degeorge et al. (1999) present evidence that U.S. firms use accounting discretion to avoid reporting small losses.[12] Thus, the incidence of small profits relative to small losses indicates the extent to which a set of firms uses accounting discretion to avoid losses. A firm-year observation is classified as small profit (small loss) if positive (negative) after-tax net income falls within the range of 1 percent of lagged total assets.[13] Our first earnings management proxy (*EM1*) is the ratio of small profits to small losses for the set of firms defined by industry and country, and by public versus private firms.

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[10] The two Directives had to be transformed into national laws by the member states during the late 1980s and early 1990s. As a result, accounting standards across EU member states are fairly similar, though not necessarily equal in every respect. Explicit transformation choices in the directives as well as so-called "soft transformations" lead to remaining differences (e.g., Stolowy and Jeny-Cazavan 2001). We address this issue in Section V by analyzing to what extent residual differences in the accounting rules affect our results.

[11] In fact, there is evidence for the U.S. that, on average, managers use their discretion in a way that makes earnings more informative (e.g., Watts and Zimmerman 1986; Subramanyam 1996). We care about the relative informativeness of earnings and its association with reporting incentives.

[12] While firms have incentives to avoid losses of any magnitude, limitations on reporting discretion make it increasingly costly to avoid larger and larger losses.

Firms can use reporting discretion to mask or misstate economic performance. For instance, firms can overstate reported earnings to achieve certain earnings targets or report extraordinary performance in specific instances, such as an equity issuance (e.g., Teoh et al. 1998a). Similarly, in years of poor performance, firms can boost their earnings using reserves or aggressive revenue recognition practices. Common to these examples is that earnings are temporarily inflated due to accrual choices, but cash flows are unaffected. Thus, our second proxy (*EM2*) is the magnitude of accruals relative to the magnitude of operating cash flow.[14] *EM2* is computed as the median ratio of the absolute value of total accruals scaled by the corresponding value of cash flow from operations for an industry within a country, where the scaling controls for differences in firm size and performance.[15]

Controlling owners and managers can conceal changes in their firm's economic performance by smoothing reported earnings. Our third measure (*EM3*) attempts to capture the extent to which corporate insiders reduce the variability of reported earnings using accruals. It is computed as the ratio of the standard deviation of operating income divided by the standard deviation of cash flow from operations, multiplied by –1 so that higher values correspond to more earnings smoothing. Due to data restrictions, we calculate the standard deviations in the cross-section (see also, Lang et al. 2003; Lang et al. 2006).

Our final measure examines accrual choices in response to shocks to the firm's economic performance. Firms can use accruals to hide bad current performance or to underreport good current performance, effectively saving for the future. This behavior induces a negative correlation between changes in accruals and the shocks to operating cash flows. While a negative correlation is a "natural" result of accrual accounting (e.g., Dechow 1994), larger magnitudes of this negative correlation indicate, *ceteris paribus*, smoothing of reported earnings that does not reflect a firm's underlying economic performance (Myers et al. 2006).[16] Our fourth earnings management measure (*EM4*) is the contemporaneous Spearman correlation between changes in total accruals and changes in the cash flow from operations (both scaled by lagged total assets) calculated for each industry-country unit of analysis, again multiplied by –1 so that higher values indicate higher levels of earnings management.

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[13] Alternative cutoff values of 0.5 percent or 2 percent of lagged total assets yield results that are qualitatively very similar to those reported in the tables below.

[14] The exercise of reporting discretion does not always imply that earnings become less informative, but prior studies show that extreme accruals are indicative of poor earnings quality (e.g., Sloan 1996). Note further that we assume that cash flows are free of manipulation, although this is not always the case (e.g., Roychowdhury 2005).

[15] Cash flow from operations is calculated using the balance-sheet approach because U.S. style cash flow statements are generally not available for our sample of private and public European companies. Following Dechow et al. (1995), we compute the accrual component of earnings as (Δtotal current assets – Δcash) – (Δtotal current liabilities – Δshort-term debt) – depreciation expense, where Δ denotes the change over the last fiscal year. If a firm does not report information on cash or short-term debt, then the changes in both variables are assumed to be zero. We scale all accounting items by lagged total assets to ensure comparability across firms.

[16] As accounting systems likely under-react to economic shocks, insiders signaling firm performance use accruals in a way that on average results in a less negative (and in specific cases even positive) correlation with cash flows.

To mitigate potential measurement error in the individual scores, we base most of our analyses on an aggregate measure of earnings management. Specifically, we transform each individual score into percentage ranks (ranging from 0 to 100) and combine the ranks by averaging into an aggregate index of earnings management, denoted EM_{aggr}.[17]

Data, Sample Selection, and Descriptive Statistics

The primary source of financial data is the June 2005 version of the Amadeus Top 200,000 database supplied by Bureau van Dijk. Amadeus provides financial statement data for a vast set of European private and public companies and is compiled from several well-established national information collectors. Since its coverage is less detailed in the early years, we focus on the seven-year period from 1997 to 2003. Amadeus provides consolidated financial statements when they are available and parent-only accounts otherwise. Thus, our analysis is based on a firm's primary set of financial statements from an informational perspective. The main advantage of the relatively new Amadeus database is that it includes privately held corporations, allowing us to focus on an economically important group of firms that is relatively under-represented in academic research.

The initial sample consists of all firm-year observations from private and public companies that have their domicile in one of the 15 member states of the EU by the year 2003 and for which current year's net income and previous year's total assets are available on Amadeus. By adopting size restrictions similar to those laid out in the Fourth EU Directive, we explicitly exclude small privately held firms to which the EU directives may not fully apply.[18] Specifically, we require privately held firms to meet at least two of the following three criteria in every year: (1) total assets greater than EUR 2.5 million, (2) sales greater than EUR 5 million, and (3) number of employees greater than 50. We also exclude banks, insurance companies and other financial holdings (SIC codes 6000–6799), public administrative institutions (SIC codes above 9000), as well as privately held subsidiaries of quoted companies as indicated in Amadeus. Investment, financing, and operating decisions in the latter firm category are likely to be influenced by parent companies, which could bias results if they were included in the analyses.

Next, we screen out private firms whose legal form is not equal to the status of corporations or its national equivalents. Strictly speaking, the argument that private and public firms face similar accounting rules only applies to corporations and does not extend to other legal forms like sole proprietorships or partnerships. We further eliminate all firms that go public during the sample period because prior work shows that those firms are subject to systematically higher levels of earnings management (e.g., Teoh et al. 1998b).[19] Finally, we also discard observations from two sample countries with missing accounting and legal institutional data. Amadeus does not provide data on operating income and depreciation expenses for companies from Ireland, and several institutional proxies used in the analysis are missing for Luxembourg.[20]

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[17] See Section IV for an assessment of the sensitivity of the results to alternative dependent variables (Table 5, Panel B).

[18] The Fourth EU Directive distinguishes between small-, medium-, and large-sized companies depending on three criteria: balance sheet total, net turnover, and average number of employees (Articles 11 and 27). Small- and medium-sized companies are subject to certain exemptions from reporting requirements, e.g., they are allowed to draw up abridged balance sheets and income statements.

[19] Amadeus does not provide time-series data on a firm's listing status and classifies each corporation by its most recent status available (i.e., either listed or not). However, by comparing firms' listing status across two annual releases of the Amadeus database we were able to identify formerly private companies that went public over the sample period. We thank Leora Klapper for providing this information. In addition, Amadeus indicates for a small subset of firms the date of the IPO, which also lets us identify firms with changes in their listing status. We use both sources of information to eliminate firms with recent IPOs. The results are very similar if we keep those recently listed firms in our analyses. In related tests (not reported), we confirm that our results are not unduly affected by quoted firms that are cross-listed on multiple international exchanges.

To mitigate the influence of outliers and potential data errors we truncate accounting items needed in the calculation of our proxies at the 1st and 99th percentile and delete firm-year observations where accounting items are exactly equal to zero, most likely indicating missing data. For robustness, we check that our results do not hinge on either of these two design choices. The final sample consists of 378,122 firm-year observations from private and publicly traded, nonfinancial corporations over the fiscal years 1997 to 2003 across 13 European countries.

Several of our earnings management proxies have to be computed for a group of firms. Prior work typically uses *country*-level observations (e.g., Leuz et al. 2003). To better control for firm characteristics, we choose a finer partition and define a unit of analysis within each country on the *industry*-level using the industry classification in Campbell (1996). That is, we calculate the individual and aggregate earnings management scores by country and industry for both public and private firms, resulting in 312 possible observations (= 13 countries × 12 industry classes × 2 firm types).[21] We require a minimum of ten firm-year observations per unit of analysis. This requirement reduces the sample to a total number of 269 industry-level observations, of which 151 are from privately held firms and 118 are from publicly listed firms.[22]

Table 1 presents descriptive statistics for the four individual earnings management measures (*EM1* through *EM4*) as well as the overall earnings management index (EM_{aggr}). On the country-level, mean values of industry observations from privately held companies generally exceed their public counterparts. For EM_{aggr}, only Greece exhibits less earnings management among private companies, which may partly reflect the fact that Greek publicly traded firms exhibit more earnings management than public companies in any other sample country.[23] On the other end of the spectrum, public firms from the U.K. and Finland exhibit low levels of earnings management. For the sample as a whole, mean and median values calculated from public companies are significantly lower than private company means and medians for both the individual and aggregate EM measures, except that median *EM4* is not significantly different between the two groups.

Table 2 presents descriptive statistics for firm characteristics used as control variables in the multivariate tests. We choose proxies for which prior work suggests an association with the level of earnings management (or accruals) and which are also likely to capture the apparent heterogeneity across private and public firms. Firm size (*SIZE*) is measured as the book value of total assets at the end of the fiscal year (in EUR thousands). Since access to capital and other corporate financing decisions depend on the extent of agency costs and asymmetric information (e.g., Titman and Wessels 1988; Rajan and Zingales 1995) and

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[20] If we include Luxembourg in the analysis adopting the legal institutional data from Belgium, all the results and inferences remain the same.

[21] Observations with missing industry data in Amadeus are grouped together in a separate industry class. If we delete this *ad hoc* group from our analyses, the results and the inferences remain unchanged.

[22] As we increase the required number of firm-year observations per group to n ≥ 25 (and then n ≥ 50), we lose 41 (70) public firm observations, and 3 (7) private firm observations, respectively. We replicated subsequent analyses using these two larger minimum numbers of observations. The substantial reduction in public firm observations weakens the statistical power, resulting in lower significance levels for the public firm indicator in some of the analyses but without changing the tenor of the results.

[23] This finding is consistent with Leuz et al. (2003) where Greece (together with Austria) ranked highest in terms of earnings management for an international sample of quoted companies.

TABLE 1
Descriptive Statistics for the Individual and Aggregate Earnings Management Scores by Country and Listing Status

Country	Listing Status	Firm-Years	Industry Obs.	*EM1*	*EM2*	*EM3*	*EM4*	EM_{aggr}
Austria	private	1,944	11	4.098	0.801	−0.461	0.931	75.1
	public	20	1	0.000	0.729	−0.205	0.983	70.3
Belgium	private	27,702	12	3.055	0.744	−0.542	0.901	65.9
	public	204	9	0.722	0.620	−0.595	0.868	43.2
Denmark	private	9,400	12	2.902	0.641	−0.621	0.833	45.7
	public	272	11	0.818	0.601	−0.699	0.772	35.3
Finland	private	9,334	11	3.150	0.521	−0.703	0.812	30.9
	public	354	10	1.000	0.430	−0.747	0.700	24.0
France	private	39,330	12	3.764	0.695	−0.619	0.873	57.3
	public	1,041	11	3.245	0.505	−0.791	0.814	28.6
Germany	private	8,271	12	4.613	0.796	−0.474	0.933	80.4
	public	1,159	10	2.792	0.614	−0.584	0.906	52.8
Greece	private	8,191	11	1.574	0.728	−0.600	0.893	50.0
	public	742	11	1.945	0.608	−0.512	0.936	55.8
Italy	private	71,081	12	3.984	0.733	−0.496	0.935	78.7
	public	327	10	2.700	0.676	−0.576	0.913	59.3
The Netherlands	private	21,679	12	1.830	0.609	−0.634	0.870	40.2
	public	570	10	2.333	0.503	−0.728	0.856	34.1
Portugal	private	9,062	10	5.881	0.744	−0.504	0.937	79.3
	public	140	6	4.000	0.823	−0.624	0.885	64.2
Spain	private	61,021	12	4.457	0.665	−0.573	0.889	61.6
	public	470	12	1.667	0.422	−0.587	0.852	41.4
Sweden	private	22,554	12	2.386	0.632	−0.630	0.834	40.3
	public	230	6	0.700	0.687	−0.811	0.771	37.6
United Kingdom	private	79,051	12	1.992	0.686	−0.586	0.858	49.0
	public	3,973	11	2.206	0.466	−0.706	0.750	24.2

(*continued on next page*)

TABLE 1 (Continued)

Mean	private	368,620	151	3.335	0.691	−0.573	0.884	57.9
(Total)	public	9,502	118	1.977**	0.565**	−0.655**	0.836**	41.1**
Median	private		151	2.882	0.695	−0.575	0.895	57.1
	public		118	1.500**	0.542**	−0.649**	0.879	40.1**
Standard Deviation	private		151	2.260	0.107	0.091	0.059	18.9
	public		118	1.964	0.159	0.198	0.144	19.7

** Indicates statistical significance at the 1 percent level (two-tailed).

The base sample consists of 378,122 firm-year observations from private and public, nonfinancial companies over the fiscal years 1997 to 2003 across 13 countries of the European Union. We exclude Ireland and Luxembourg because of insufficient financial and institutional data. We also exclude firms that change their listing status during the sample period. Financial data are obtained from the June 2005 version of the Amadeus Top 200,000 database. The computations are at the industry-level using the classification in Campbell (1996), i.e., we calculate each score by country, industry, and listing status. We require a minimum of ten firm-year observations per subgroup. We truncate firm-level realizations at the first and 99th percentile before computing the score per subgroup. The table presents mean values by country and listing status for the four individual earnings management scores and the aggregate earnings management index (see Leuz et al. 2003). *EM1* is the number of "small profits" divided by the number of "small losses." A firm-year observation is classified as small profit (small loss) if positive (negative) net income falls within the range of 1 percent of lagged total assets. *EM2* is the median ratio of the absolute value of total accruals to the absolute value of cash flow from operations. Total accruals are calculated as follows: (Δ total current assets − Δ cash) − (Δ total current liabilities − Δ short-term debt) − depreciation expense. Cash flow from operations is equal to operating income minus total accruals. *EM3* is the ratio of the cross-sectional standard deviations of operating income and cash flow from operations (multiplied by −1). *EM4* is the Spearman correlation between the change in total accruals and the change in cash flow from operations (multiplied by −1). All accounting items are scaled by lagged total assets. The aggregate earnings management index, EM_{aggr}, is the average percentage rank across all four individual scores, *EM1* to *EM4*. EM scores are constructed such that higher values imply higher levels of earnings management. We evaluate differences between sample means (medians) using t-tests (Wilcoxon rank sum tests).

TABLE 2
Descriptive Statistics for the Firm-Level Control Variables by Country and Listing Status

Country	Listing Status	Industry Obs.	*SIZE*	*LEV*	*GROWTH*	*ROA*	*CYCLE*	*AUDIT*	*AGE*	*OWNER*
Austria	private	11	27,900	46.2%	4.2%	3.0%	94.6	NA	14.0	100.0%
	public	1	182,845	39.9%	11.8%	3.6%	137.5	100.0%	13.5	46.2%
Belgium	private	12	11,676	23.8%	5.5%	2.2%	90.3	25.9%	19.5	98.5%
	public	9	283,906	31.7%	4.9%	4.2%	126.0	85.2%	74.5	41.7%
Denmark	private	12	14,462	27.4%	4.5%	3.7%	71.9	76.0%	14.5	100.0%
	public	11	188,272	27.8%	6.6%	4.3%	88.2	99.5%	40.5	29.0%
Finland	private	11	10,014	26.2%	7.7%	5.7%	62.4	95.5%	13.0	100.0%
	public	10	191,941	30.3%	7.8%	6.7%	89.5	99.8%	43.3	31.3%
France	private	12	12,993	20.8%	6.7%	2.9%	81.8	14.8%	20.5	79.5%
	public	11	161,900	9.0%	4.9%	4.0%	105.7	43.8%	45.0	51.0%
Germany	private	12	38,195	52.6%	3.4%	2.4%	71.1	NA	16.0	100.0%
	public	10	211,650	50.4%	3.2%	3.0%	103.8	91.7%	66.5	66.4%
Greece	private	11	13,202	8.7%	11.3%	3.1%	151.0	0.0%	12.0	52.6%
	public	11	58,989	6.7%	10.6%	5.3%	198.4	35.4%	22.0	34.8%
Italy	private	12	21,705	37.4%	7.1%	1.1%	122.7	0.0%	19.5	81.0%
	public	10	406,272	36.7%	4.1%	3.6%	142.4	98.6%	59.8	50.8%
The Netherlands	private	12	20,433	30.1%	5.4%	4.1%	84.5	87.4%	20.5	100.0%
	public	10	571,746	37.1%	5.9%	6.4%	94.5	100.0%	66.0	16.4%
Portugal	private	10	13,921	21.4%	7.8%	2.2%	126.0	0.0%	20.5	65.0%
	public	6	69,293	28.0%	5.4%	1.6%	188.7	95.1%	39.5	40.3%
Spain	private	12	14,150	17.1%	10.7%	4.0%	111.4	58.1%	15.5	85.8%
	public	12	293,161	22.6%	9.6%	5.6%	148.7	99.4%	42.0	26.7%
Sweden	private	12	11,224	37.0%	7.4%	3.2%	67.6	NA	18.5	100.0%
	public	6	80,994	42.8%	6.1%	2.9%	87.8	100.0%	32.3	32.5%
United Kingdom	private	12	18,628	17.3%	6.5%	3.8%	60.7	80.7%	18.0	100.0%
	public	11	182,965	23.5%	7.3%	5.3%	75.4	98.2%	33.0	13.9%

(*continued on next page*)

TABLE 2 (Continued)

Mean	private	151	25,802	29.4%	7.0%	3.1%	93.4	47.0%	17.2	88.9%
	public	118	475,551**	29.2%	7.6%	4.5%**	116.0**	79.7%**	46.3**	36.7%**
Median	private	151	15,767	27.7%	6.6%	3.0%	84.4	52.8%	17.0	99.9%
	public	118	178,307**	29.2%	6.3%	4.3%**	106.1**	97.3%**	41.0**	33.9%**
Standard Deviation	private	151	38,101	14.0%	3.7%	1.5%	37.7	39.0%	5.2	16.4%
	public	118	1,181,688	14.9%	5.1%	2.3%	49.2	31.3%	26.2	18.1%

** Indicates statistical significance at the 1 percent level (two-tailed).

The base sample consists of 378,122 firm-year observations from private and public, nonfinancial companies over the fiscal years 1997 to 2003 across 13 countries of the European Union. We exclude Ireland and Luxembourg because of insufficient financial and institutional data. We also exclude firms that change their listing status during the sample period. Financial, auditor, and ownership data are obtained from the June 2005 version of the Amadeus Top 200,000 database. The analysis is based on country-industry medians by listing status, i.e., private and public firms in a particular industry and country form separate subgroups. We require a minimum of ten firm-year observations per subgroup. We truncate firm-level realizations at the first and 99th percentile before computing the subgroup median. The industry classification is based on Campbell (1996). The table presents medians for the control variables by country and listing status. *SIZE* is the book value of total assets at the end of the fiscal year (in EUR thousands). Financial leverage *(LEV)* is measured as the ratio of noncurrent liabilities to the sum of noncurrent liabilities plus book value of equity. *GROWTH* is the annual percentage change in revenue. *ROA* stands for yearly return on assets and equals net income divided by lagged total assets. *CYCLE* represents the operating cycle (in days) calculated as (yearly average accounts receivable)/(total revenue/360) + (yearly average inventory)/(cost of goods sold/360). Cost of goods sold is equal to total revenue minus operating income. *AUDIT* is the weighted market share of Big 5 audit firms and calculated as Big 5 clients' total assets divided by all firms' total assets under audit. *AGE* is firm age in years. Ownership concentration *(OWNER)* is measured as the percentage of direct holdings by the single largest shareholder of the firm. We evaluate differences between sample means (medians) of private and public firms using t-tests (Wilcoxon rank sum tests).

these attributes likely differ between public and private companies, we include financial leverage as a control variable. We calculate *LEV* as the ratio of total non-current liabilities to total assets.

Other potential sources of variation in accruals and between privately held and public companies are firm growth, profitability, the length of the operating cycle, audit quality, firm age, and ownership concentration. *GROWTH* is defined as the annual percentage change in revenue. Profitability is measured as return on assets (*ROA*) defined as net income divided by lagged total assets. Following Dechow (1994), we compute the length of the operating cycle in days, *CYCLE*, as (yearly average accounts receivable)/(total revenue /360) + (yearly average inventory)/(cost of goods sold/360). Since the cost of goods sold number is not reported separately for many companies, we use total revenue minus operating income instead. We proxy for audit quality, *AUDIT*, by the weighted market share of Big 5 audit firms, calculated as Big 5 clients' total assets divided by all firms' total assets under audit.[24] *AGE* is the number of years since incorporation. Finally, we measure ownership concentration, *OWNER*, as the percentage of direct holdings by the single largest shareholder as provided in Amadeus. All firm attributes are computed as industry-level medians.

Table 2 shows that, as expected, public firms are on average larger, more profitable, have longer operating cycles, rely more frequently on high-quality auditors, exist for a longer time period, and have more dispersed ownership than private firms. With respect to financial leverage and sales growth, no clear pattern arises.[25] Table 3 provides descriptive country-level information (raw values and binary classifications) on the legal, institutional, and capital market variables discussed and used in the next sections.

IV. THE ROLE OF PUBLIC EQUITY MARKETS AND LEGAL ENFORCEMENT

We begin our analysis of firms' reporting incentives by studying the effect of public equity markets and legal enforcement—an institutional variable that we expect to influence the reporting behavior of both private and public firms. We use a binary variable to indicate observations stemming from firms with publicly traded equity securities (*PUBL*). The quality of legal enforcement (*LEGAL*) is measured by the average score across three proxies from La Porta et al. (1998): (1) an index of the judicial system's efficiency, (2) an index of the rule of law, and (3) the level of corruption. *LEGAL* ranges from 0 to 10 with higher values corresponding to stricter legal enforcement (see Table 3).[26]

In selected analyses, we also examine the effects of a country's legal origin (i.e., English, Scandinavian, French, or German), which is probably the most exogenous classification of international legal institutions. However, the arguments in La Porta et al. (1998) and Ball et al. (2000), as well as results from a cluster analysis in Leuz et al. (2003, Table 3, Panel B) suggest that legal origins are more appropriately viewed as a summary measure

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[24] We only can collect audit information for about 65 percent of the firm-year observations. In particular, auditor data for private firms in Austria, Germany, and Sweden does not exist.

[25] The insignificant mean and median differences for *LEV* may be due to our inability to distinguish between various forms of debt. Amadeus does not provide detailed data on firms' liabilities on a consistent basis across countries, which prevents us from using leverage proxies that are specific to particular forms of debt. Untabulated sensitivity analyses, however, show that our results do not hinge on the specific definition of financial leverage.

[26] Note that the variables from La Porta et al. (1998) used here and several other institutional variables introduced in Section V were measured before the start of our sample period and therefore are not contemporaneous with our other variables. However, we do not believe this has a major effect on our analysis because we rely only on relative values of these variables and because countries' institutional features tend to change only slowly over time.

TABLE 3
Descriptive Statistics for the Institutional Variables by Country

	Legal and Enforcement Quality		Tax Alignment		Accrual Accounting Rules	Securities Regulation and Investor Protection		Capital Market Structure	
Country	***ORIGIN***	***LEGAL***	***TAX***	***RATE***	***ACCRUAL***	***SECREG***	***ANTIDIR***	***SYSTEM***	***DEV***
Austria	German	9.36 (1)	1	34.0% (1)	0.55 (0)	0.18 (0)	2 (0)	−3.55 (0)	5.23 (1)
Belgium	French	9.44 (1)	1	40.2% (1)	0.64 (1)	0.34 (0)	0 (0)	−2.27 (1)	4.33 (0)
Denmark	Scandinavian	10.00 (1)	0	32.0% (0)	0.55 (0)	0.50 (1)	2 (0)	−2.81 (0)	4.69 (0)
Finland	Scandinavian	10.00 (1)	1	28.0% (0)	0.77 (1)	0.49 (0)	3 (0)	−3.10 (0)	4.99 (0)
France	French	8.68 (0)	1	40.0% (1)	0.64 (1)	0.58 (1)	3 (0)	−2.83 (0)	6.01 (1)
Germany	German	9.05 (0)	1	53.0% (1)	0.41 (0)	0.21 (0)	1 (0)	−1.64 (1)	7.26 (1)
Greece	French	6.82 (0)	1	40.0% (1)	0.41 (0)	0.38 (0)	2 (0)	−4.48 (0)	2.58 (0)
Italy	French	7.07 (0)	1	37.0% (1)	0.59 (0)	0.46 (0)	1 (0)	−2.79 (1)	5.01 (0)
The Netherlands	French	10.00 (1)	0	35.0% (0)	0.77 (1)	0.62 (1)	2 (0)	−1.65 (1)	7.31 (1)
Portugal	French	7.19 (0)	1	37.4% (1)	0.55 (0)	0.55 (1)	3 (0)	−4.26 (0)	4.24 (0)
Spain	French	7.14 (0)	1	35.0% (1)	0.77 (1)	0.50 (1)	4 (1)	−2.71 (1)	5.71 (1)
Sweden	Scandinavian	10.00 (1)	1	28.0% (0)	0.64 (1)	0.45 (0)	3 (0)	−1.60 (1)	6.68 (1)
United Kingdom	English	9.22 (1)	0	31.0% (0)	0.86 (1)	0.72 (1)	5 (1)	−0.76 (1)	7.14 (1)

The table presents raw and dichotomized indicator values (in parentheses) of the institutional proxies used in the analyses across the 13 sample countries from the European Union. We use institutional factors that capture countries' legal tradition and enforcement, the role of corporate taxes in determining financial accounts, firms' discretion in accounting for accruals, the strength of securities regulation, and minority shareholder protection, as well as the structure and development of organized capital markets. The legal variables consist of two measures from La Porta et al. (1998): the classification of the legal origin *(ORIGIN)*, and the quality of the legal system and enforcement *(LEGAL)* measured by the mean of three institutional variables (i.e., efficiency of the judicial system, rule of law, and corruption index). *TAX* is an indicator variable taking on the value of 1 if financial accounts for external reporting and tax purposes are highly aligned, and 0 otherwise (see Alford et al. 1993; Hung 2001). We assume a tax status of 1 for the three countries with missing tax information (Austria, Greece, and Portugal). *RATE* stands for the average corporate tax rate in percent of earnings before taxes (Source: *IMD World Competitiveness Yearbook 2000). ACCRUAL* is the accrual index from Hung (2001) (updated for European countries by Comprix et al. [2003]), and captures differences in accrual accounting rules across countries. *SECREG* captures the strength of securities regulation mandating and enforcing disclosures for publicly listed firms. It is measured as the mean of the disclosure index, the liability standard index and the public enforcement index from La Porta et al. (2006). *ANTIDIR* is the antidirector rights index from La Porta et al. (1998) capturing the legal protection of minority shareholders. The capital market variables consist of two measures from Beck and Levine (2002): *SYSTEM* captures whether a country's financial system is more market- or bank-based, and equals the natural log of (total value of stock traded/claims on the private sector by commercial banks), both scaled by GDP. Financial development, *DEV,* is a measure of the overall activity of financial intermediaries and of capital market development. It equals the natural log of (total value of stock traded*claims on the private sector by financial institutions), both scaled by GDP. For most of our analyses we transform the continuous institutional factors into binary variables splitting by the median (except for *TAX*RATE* where we use 34 percent as a cut-off, and *ANTIDIR* where we split at 4, which is commonly viewed as an indication of high investor protection; see discussion in the text for details).

that captures a multitude of institutional factors, rather than a proxy for the quality of legal enforcement. For this reason, we view these results more as a general robustness check.

Univariate Comparisons across the Capital Market and Legal Dimensions

Table 4 provides univariate tests across the two main incentive dimensions under study. Panel A summarizes the results of within-country analyses effectively holding the institutional environment constant. We test for differences between private and public firms' individual earnings management scores ($EM1$ through $EM4$) using within-country Monte Carlo simulation.[27] For the aggregate earnings management index (EM_{aggr}), we run within-country univariate regressions with an indicator variable to compare earnings management for public versus private firms. Note that these regressions are based on a small number of industry observations per country and hence have relatively low power—we provide them primarily for descriptive purposes. The evidence in Panel A suggests that earnings management is significantly more pervasive among private firms than among public firms for most, but not all, individual countries. This result holds for each individual earnings management score as well as for EM_{aggr}.

Panel B of Table 4 reports pairwise Spearman correlations. We note that all four individual earnings management measures are highly correlated and well represented by the aggregate index. Since the Amadeus database has not been used much in prior studies, we also benchmark our earnings management measures with those in Leuz et al. (2003) based on public firms from the frequently used Worldscope database. In (untabulated) analyses we find that the correlations between their proxies and our measures calculated on the country-level for public firms only are above .50 for the individual EM scores ($EM1$ produces the lowest correlation with $\rho = .52$, followed by $EM2$ with $\rho = .65$), and above .90 for the aggregate index. As reliably measuring loss avoidance requires a substantial number of firm-years, it is possible that the relatively low correlation of $EM1$ with the respective proxy in Leuz et al. (2003) is driven by the smaller number of public firms in the Amadeus database. As a robustness check, we repeat all our analyses dropping $EM1$ (or any other individual EM score) from the construction of the aggregate index. The results are very similar.[28] The correlation between the EM scores and listing status or legal enforcement quality is significantly negative, except in one case.

Panel C of Table 4 compares the aggregate earnings management index across groups of countries defined by their legal origins, which broadly capture investor protection and ownership regimes across countries (La Porta et al. 1999). The ordering of the mean and median values of EM_{aggr} shows the lowest scores for firms in countries with English and Scandinavian legal origins, followed by firms in countries with French origin. Firms with German legal origin exhibit the highest earnings management. These findings are consistent with Leuz et al. (2003). Thus, even within the set of EU countries, we observe significant variation in the level of earnings management reflecting institutional differences.

Panel D of Table 4 reports average EM_{aggr} for subgroups defined by the two main variables, quality of legal enforcement and listing status, in a simplified analysis that serves as a precursor to the more complete analysis in subsequent tables. We create a binary variable indicating high and low enforcement quality countries splitting *LEGAL* by the median. As hypothesized, publicly traded firms in countries with strict enforcement show the lowest level of earnings management. In contrast, private firms located in low-quality legal enforcement environments exhibit the highest levels of earnings management. The

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[27] For each individual EM score, we generate test statistics based on 1,000 random draws from the population of private firms and compare the actual realization stemming from all the public firms within the same country to the generated null-hypothesis distribution.

[28] Factor analysis further supports the use of an aggregate index. We find that there is only one factor with an Eigenvalue above 1 and that all of the four individual scores exhibit substantial loadings onto this factor. Moreover, when we conduct the analyses using any one of the individual measures alone, the results are generally consistent, albeit at times weaker in statistical significance, which is consistent with the notion that aggregation across measures reduces measurement error.

TABLE 4
Univariate Analysis of Earnings Management Variables, Listing Status (*PUBL*) and Legal Enforcement (*LEGAL*)

Panel A: Within-Country Analysis of Public versus Private Firms

Variables	*EM1*	*EM2*	*EM3*	*EM4*	EM_{aggr}
# higher/not distinguishable/ lower EM scores for Private Firms	(4/9/0)	(9/4/0)	(9/3/1)	(8/3/2)	(8/5/0)

Panel B: Spearman Correlation Coefficients

Variables (n = 269)	*EM1*	*EM2*	*EM3*	*EM4*	EM_{aggr}
EM2	0.325**				
EM3	0.233**	0.507**			
EM4	0.200**	0.396**	0.606**		
EM_{aggr}	0.570**	0.765**	0.808**	0.752**	
PUBL	−0.398**	−0.444**	−0.255**	−0.097	−0.381**
LEGAL	−0.241**	−0.224**	−0.325**	−0.387**	−0.415**

Panel C: Pairwise Comparison across Legal Origins

(Variable = EM_{aggr})

Legal Origins		English	Scandinavian	French	German
English	Mean	37.1			
(n = 23)	Median	35.7			
Scandinavian	Mean	(0.32)	35.9		
(n = 62)	Median	(0.06)	36.3		
French	Mean	(4.54)**	(7.31)**	54.1	
(n = 150)	Median	(3.77)**	(6.06)**	54.2	
German	Mean	(7.57)**	(10.18)**	(5.05)**	70.3
(n = 34)	Median	(5.29)**	(7.04)**	(4.23)**	72.4

Panel D: Pairwise Comparison across Legal Enforcement/Listing Status Subgroups

Legal Enforcement (Variable = EM_{aggr})		Listing Status		Test of Difference
		Public	Private	
High Enforcement	Mean	33.1	49.5	(5.34)**
Quality	Median	33.7 (n = 58)	46.5 (n = 82)	(4.70)**
Low Enforcement	Mean	48.8	67.8	(6.36)**
Quality	Median	49.2 (n = 60)	70.6 (n = 69)	(5.54)**
Test of Difference		(4.72)** (4.32)**	(6.84)** (6.04)**	

(*continued on next page*)

TABLE 4 (Continued)

** Indicates statistical significance at the 1 percent level (two-tailed).

The sample comprises 269 industry-level observations from 13 European countries. The aggregate earnings management index, EM_{aggr}, is the average percentage rank across all four individual earnings management scores, *EM1* to *EM4*, as described in Table 1. EM scores are constructed such that higher values imply higher levels of earnings management. In Panel A, we assess differences in the four individual EM scores and the aggregate EM index across public and private firms within country using Monte Carlo simulation (*EM1* to *EM4*) or regression analysis with the public/private indicator as independent variable (EM_{aggr}). The table reports the number of countries where the EM scores for private firms are (1) significantly higher, (2) not distinguishable, and (3) significantly lower than the EM scores for public firms at the 10 percent level (two-tailed). In Panel B, we report Spearman correlation coefficients for *EM1* through *EM4*, the aggregate EM index and the incentive variables of interest. *PUBL* is an indicator variable taking on the value of 1 if the observation stems from publicly traded firms, and 0 otherwise. *LEGAL* stands for the mean of three variables in La Porta et al. (1998), which measure the quality of the legal system and enforcement (i.e., efficiency of the judicial system, rule of law, and corruption index). In Panels C and D, we report means and medians of EM_{aggr} by subgroup (i.e., legal origin and legal enforcement/listing status). We assign observations to the high (low) legal enforcement quality subgroup according to the median value of *LEGAL*. In parentheses, we report t-stats and Z-stats for pairwise differences across groups.

off-diagonal comparisons show both variables appear to play an equal role in the pervasiveness of earnings management, as the private firms/high enforcement quality and the public firms/low enforcement quality cells are statistically indistinguishable. The results of this simplified analysis suggest that both incentive dimensions contribute separately to the reporting behavior of European private and public firms.

Multivariate Tests of the Reporting-Incentives Hypotheses

Table 5 presents results of regressions that examine the role of capital market forces and legal factors, separately and jointly, and include additional controls for differences in firm characteristics. In Panel A, we use the aggregate earnings management index based on the four Leuz et al. (2003) measures as the dependent variable. Panel B provides sensitivity checks using alternative proxies for earnings informativeness as dependent variables. Throughout the discussion, assessments of statistical significance are based on heteroscedasticity-corrected t-statistics.[29]

The first two columns of Table 5, Panel A, examine the effects of listing status and legal enforcement separately on the level of earnings management. We introduce controls for size, financial leverage, growth, return on assets, and operating cycle. These variables are intended to control for heterogeneity in firms' business processes, which could affect the magnitude of the accruals and the properties of reported earnings. In addition, we include industry controls based on the industry classification in Campbell (1996). The coefficient on *PUBL* in column one is significantly negative. This finding is consistent with our hypothesis that the first-order effect of public equity markets is to increase the demand for informative earnings. Results in column two for the separate effect of legal enforcement also show a significantly negative coefficient on *LEGAL*. Except for operating cycle, all the firm-level controls are statistically significant at the .05 level or better. The sign switch of the size coefficient occurs presumably because firm size captures the public versus private dimension, which is omitted in this specification.

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[29] As the legal variable varies only at the country level, we repeat our analyses using heteroscedasticity-corrected standard errors *clustered* by country, resulting in a more conservative assessment of statistical significance. However, the inferences are the same and the legal variable remains significant at the .05 level in all models presented in Table 5.

TABLE 5
Earnings Management and Reporting Incentives: The Role of Listing Status and Legal Enforcement

$$EM_{aggr,i} = \alpha_0 + \alpha_1 PUBL_i + \alpha_2 LEGAL_i + \alpha_3 SIZE_i + \alpha_4 LEV_i + \alpha_5 GROWTH_i + \alpha_6 ROA_i + \alpha_7 CYCLE_i + \Sigma\alpha_j \text{Industry Controls}_i + \varepsilon_i$$

Panel A: Aggregate Earnings Management Index as Dependent Variable[a]

	Capital Market and Legal Incentives (Base Model)			Base Model Plus Additional Controls			Balanced Panel	Largest Private Firms
Variables	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**	**Model 7**	**Model 8**
n	269	269	269	234	269	269	110	264
PUBL	−23.568**	—	−18.997**	−14.052**	−18.590**	−16.393**	−24.929**	−16.353**
	(−6.18)		(−4.86)	(−3.56)	(−4.54)	(−3.71)	(−3.82)	(−6.41)
LEGAL	—	−7.614**	−6.270**	−5.015**	−6.273**	−6.357**	−7.528**	−5.669**
		(−6.32)	(−5.22)	(−4.12)	(−5.24)	(−5.30)	(−3.69)	(−4.66)
Intercept	10.293	159.055**	90.543**	93.435**	90.234**	84.471**	82.947**	94.205**
	(0.81)	(9.83)	(4.56)	(4.79)	(4.57)	(3.99)	(2.77)	(4.55)
SIZE	2.737*	−3.965**	1.457	0.714	1.530	1.573	2.712	1.033
	(2.16)	(−5.45)	(1.08)	(0.52)	(1.12)	(1.16)	(1.26)	(0.76)
LEV	21.016*	43.750**	30.999**	21.075*	31.124**	29.870**	20.832	15.355#
	(2.38)	(5.03)	(3.49)	(2.11)	(3.49)	(3.35)	(1.35)	(1.70)
GROWTH	−25.056	−58.373*	−62.154*	−49.938#	−63.791*	−57.262*	−98.769	−55.758#
	(−1.00)	(−2.14)	(−2.33)	(−1.82)	(−2.37)	(−2.11)	(−1.57)	(−1.69)
ROA	−3.370**	−2.907**	−2.624**	−2.634**	−2.623**	−2.635**	−1.473	−1.782**
	(−4.92)	(−4.25)	(−3.96)	(−3.99)	(−3.95)	(−4.02)	(−1.18)	(−2.73)
CYCLE	0.214**	0.001	0.074#	0.078#	0.074#	0.078#	0.085	0.055
	(6.58)	(0.02)	(1.74)	(1.79)	(1.73)	(1.82)	(1.41)	(1.36)
AUDIT	—	—	—	−5.224	—	—	—	—
				(−1.44)				
AGE	—	—	—	—	−0.020	—	—	—
					(−0.35)			
OWNER	—	—	—	—	—	0.058	—	—
						(1.06)		
Industry Controls	included	included	included	included	included	included	included	included
R^2	54.7%	55.2%	59.6%	60.3%	59.6%	59.8%	61.8%	54.8%

(continued on next page)

TABLE 5 (Continued)

Panel B: Sensitivity Analyses with Alternative Dependent Variables[b]

Variables	Earnings Conservatism Index EC_{aggr} (higher values = less earnings conservatism)	Alternate Earnings Management Index EM_{alt} (higher values = more earnings management)	Overall Earnings Informativeness Index EI_{aggr} (higher values = less informative earnings)
n	252	258	249
PUBL	−16.197** (−3.23)	−10.619** (−2.73)	−17.074** (−5.11)
LEGAL	−3.456* (−2.35)	−4.738** (−3.88)	−4.771** (−4.43)
Intercept	92.120** (3.25)	71.365** (3.13)	79.038** (4.33)
SIZE	−2.557 (−1.37)	0.589 (0.38)	0.632 (0.52)
LEV	36.740** (4.41)	27.607** (3.63)	28.415** (4.69)
GROWTH	−17.332 (−0.48)	−11.785 (−0.44)	−26.961 (−1.13)
ROA	2.216** (3.44)	−0.967# (−1.79)	−0.822# (−1.75)
CYCLE	0.068 (1.28)	0.080* (2.02)	0.067# (1.74)
Industry Controls	included	included	included
R^2	35.1%	43.9%	59.2%

(*continued on next page*)

TABLE 5 (Continued)

**, *, and # Indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels (two-tailed), respectively.

[a] In Panel A, the sample comprises a maximum of 269 industry-level observations from 13 European countries. The dependent variable, EM_{aggr}, is the average percentage rank across all four individual earnings management scores, *EM1* to *EM4*, as described in Table 1. EM scores are constructed such that higher values imply higher levels of earnings management. *PUBL* is an indicator variable taking on the value of 1 if the observation stems from publicly traded firms, and 0 otherwise. *LEGAL* stands for the mean of three variables in La Porta et al. (1998), which measure the quality of the legal system and enforcement (i.e., efficiency of the judicial system, rule of law, and corruption index). *SIZE* is the book value of total assets at the end of the fiscal year (in EUR thousands). We use the natural log of the size variable in the analysis. Financial leverage *(LEV)* is measured as the ratio of noncurrent liabilities to the sum of noncurrent liabilities plus book value of equity. *GROWTH* is the annual percentage change in revenue. *ROA* stands for the yearly percentage return on assets. *CYCLE* represents the operating cycle (in days) calculated as (yearly average accounts receivable)/(total revenue/360) + (yearly average inventory)/(cost of goods sold/360). Cost of goods sold is equal to total revenue minus operating income. Industry controls based on the classification in Campbell (1996) are included but not reported. In Models 4 to 6 we further include one of the following control variables: *AUDIT* is the weighted market share of Big 5 audit firms and calculated as Big 5 clients' total assets divided by all firms' total assets under audit. *AGE* is firm age in years. Ownership concentration *(OWNER)* is measured as the percentage of direct holdings by the single largest shareholder of the firm. Model 7 presents a balanced panel including only observations from the same set of industries across countries and private and public firms (5 industries from 11 countries). In Model 8 we limit the subgroup of private firms to the largest 25 percent (based on total assets), which makes private and public firms more comparable in size. The table reports OLS coefficient estimates and t-statistics based on heteroscedasticity corrected standard errors (in parentheses).

[b] In Panel B, we present results corresponding to Model 3 in Panel A substituting the following alternative dependent variables: The earnings conservatism index, EC_{aggr}, consists of the average percentage rank across three individual conservatism measures: (1) the number of "large profits" divided by the number of "large losses" with firm-year observations classified as large if net income falls outside the range of 20 percent of lagged total assets, (2) the skewness of earnings over the middle part of the distribution (i.e., net income +/− 20 percent of lagged total assets), and (3) the coefficient on lagged negative changes in net income from a Basu (1997)-style regression as implemented by Ball and Shivakumar (2005). We also construct an alternative aggregate earnings management index, EM_{alt}, using a set of individual earnings management scores not contained in our main proxy EM_{aggr}. We include (1) the variance of the absolute value of changes in annual net income (multiplied by −1), (2) the variance of the absolute value of changes in operating income divided by the same metric for cash flow from operations (multiplied by −1), and (3) the absolute value of Jones (1991)-model discretionary accruals calculated separately by industry. Finally, we construct an overall earnings informativeness index, EI_{aggr}, combining all individual earnings management and conservatism scores (i.e., $EI_{aggr} = EM_{aggr} + EC_{aggr} + EM_{alt}$). All earnings scores are calculated as percentage ranks and constructed such that higher values imply higher levels of earnings management (lower levels of earnings conservatism). The regressions include the two incentive variables, listing status and legal enforcement, and the full set of controls (see Panel A for a description). The table reports OLS coefficient estimates and t-statistics based on heteroscedasticity corrected standard errors (in parentheses).

In column three we estimate the combined model that simultaneously includes the indicator for public firms and the legal enforcement variable. Again, we find that listing status and the quality of legal enforcement are significant, and each remains negatively related to EM_{aggr}. Together with the controls, these variables explain about 60 percent of the total variation in earnings management.

Next, we address several concerns about correlated omitted variables by extending our Model 3 with proxies for audit quality, firm age, and ownership concentration. First, it is possible that our findings partly reflect differences in audit quality across private and public firms. For instance, private firms may hire lower quality auditors because communicating via reported earnings is less important to them. Alternatively, auditors of publicly traded firms may face stronger pressures to be tough on their clients than auditors of private firms. Both of these effects would be consistent with our hypothesis in Section II, which does not distinguish between the *ex ante* screening and the *ex post* incentive effect of capital markets. Our analysis aims to shed light on the first-order effect of public markets on earnings informativeness, but it does not attempt to identify the exact mechanism(s).

Nevertheless, we check whether our results are robust to controls for audit quality. Results in column four of Table 5, Panel A show a negative, but insignificant coefficient on *AUDIT*. However, introducing this additional variable does not change the results or inferences for the two main variables, *PUBL* and *LEGAL*, suggesting that audit quality is unlikely to be the primary mechanism through which public markets reduce earnings management.

Second, we introduce firm age to control for the fact that listed firms are often more mature, which could lead to different earnings properties and in turn influence our findings. However, results in column five show that *AGE* is not significant and that, compared to Model 3, the magnitudes and significance levels of the coefficients on *PUBL* and *LEGAL* are nearly unchanged.

Third, going public usually goes hand in hand with an increase in ownership dispersion. Thus, ownership concentration is potentially an important omitted variable. Furthermore, ownership concentration is expected to have a similar effect on reporting quality among the subset of public firms as the private versus public distinction in our analysis.[30] Results for Model 6 show that *OWNER* is also not significant and the coefficients on *PUBL* and *LEGAL* remain negative and highly significant.[31]

The final two columns in Table 5, Panel A report results for alternative sample compositions. First, we construct a balanced panel where the industry composition is constant across countries and across public and private firms. To the extent that economic differences between industries across countries and listing status are systematically related to the informativeness of reported earnings, it is possible that industry composition accounts for some of the reported results. The balanced panel addresses this concern. The reduced sample includes five industries from 11 countries, resulting in a total of 110 public- and private-firm observations. The results are reported in column seven and indicate that our findings do not appear to be driven by differences in industry composition.

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[30] We confirm this conjecture by performing an analysis for the subset of public firms. When we introduce a binary indicator variable into the base model using different cut-off values of the *OWNER* distribution, we always find a negative and significant relation. This result is consistent with the notion that high ownership concentration has a negative effect on earnings informativeness. See also Fan and Wong (2002), Haw et al. (2004), and Wysocki (2004).

[31] Our results continue to hold if we include the three additional firm-level controls simultaneously in the model. Moreover, in unreported analyses we confirm that country-level factors like macroeconomic growth and per capita wealth do not unduly affect our results.

Second, as public firms are on average much larger than private firms, we also compute the EM scores using only the largest quartile of private firms in order to improve the size comparability between the two groups. Column eight shows that the results for our key variables of interest, *PUBL* and *LEGAL*, are essentially unchanged.[32]

Alternative Proxies for Earnings Informativeness

Panel B of Table 5 provides sensitivity analyses using alternative proxies for earnings informativeness. For brevity, we only report results for the specification corresponding to Model 3 in Panel A, jointly including the two incentive variables and the full set of controls. First, we examine the effect of market forces and legal enforcement on accounting conservatism, a frequently used dimension of earnings quality (see e.g., Ball et al. 2000; Ball and Shivakumar 2005; Bushman and Piotroski 2006). We construct an earnings conservatism index, EC_{aggr}, consisting of the following three measures: (1) the ratio of large profits to large losses, with large defined as net income exceeding 20 percent of lagged total assets. Smaller values of this ratio indicate a higher propensity of big bath accounting. (2) The skewness of net income deflated by lagged total assets over the middle part of the distribution (i.e., net income +/− 20 percent of lagged total assets). A more negatively skewed distribution indicates that firms recognize losses in a more timely manner. (3) We run a Basu (1997)-style regression as implemented by Ball and Shivakumar (2005) of changes in earnings on lagged changes in earnings and use the coefficient estimate of lagged negative changes in net income as a proxy for timely loss recognition.[33] Column one of Panel B reports results regressing EC_{aggr} on listing status and legal enforcement quality (plus controls). Both incentive variables are negative and significant, indicating more earnings conservatism for public firms and in countries with stricter legal enforcement.

Next, we present results using an alternative set of earnings management scores, EM_{alt}, not contained in our aggregate earnings management index. We combine the following three additional measures suggested by Lang et al. (2003) and Lang et al. (2006):[34] (1) The variance of the absolute values of changes in annual net income (multiplied by −1) as a measure of earnings variability. (2) The ratio of the variance of absolute changes in operating income to the same metric for cash flow from operations (multiplied by −1). These first two measures are intended to capture earnings smoothing. (3) The absolute value of discretionary accruals derived from a cross-sectional version of the Jones (1991) model, calculated separately for each industry. All else equal, a greater level of discretionary accruals is generally viewed as indicative of more earnings management (see Healy and Wahlen 1999). Column two of Table 5, Panel B shows that both incentive variables remain negative and highly significant using this separate set of earnings management scores as the dependent variable.

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[32] In order to explicitly account for differences in the availability of input data across countries and industry groups, we also re-estimate the regressions in Table 5, Panel A, using weighted least squares. The square root of the number of firm-level observations serves as the weighting variable. The results are very similar to those reported in the table.

[33] In separate tests (not reported), we replicate the analysis of Ball and Shivakumar (2005, Table 3) for the U.K. and extend it to the other EU countries. We find confirmatory evidence for Denmark, France, Germany, The Netherlands, and Sweden. For the remaining countries, the results are either not significant or go in the other direction. However, one problem with such an analysis in a cross-country setting (e.g., Peek et al. 2006) is that lagged earnings changes in the Basu-style regression are likely to reflect institutional differences and hence may not be a good proxy for shocks to economic earnings.

[34] To be consistent with the rest of our analyses, we do not use firm-level controls in the calculation of the individual earnings management scores (see Lang et al. 2003), but control for firm-level attributes when regressing aggregate *EM* scores on the incentive variables.

In the final column we present results using an overall earnings informativeness index, EI_{aggr}, consisting of all the individual earnings management and conservatism scores (i.e., $EI_{aggr} = EM_{aggr} + EC_{aggr} + EM_{alt}$). Higher values of EI_{aggr} indicate less informative earnings. The results are again consistent with our earlier findings.

Taken as a whole, the results throughout Table 5 are consistent with our main hypothesis that public equity markets and strong legal enforcement provide incentives to report earnings that reflect economic performance. These results hold after controlling for a variety of firm-specific factors, after substantial modifications to the sample composition, and across a number of proxies for earnings management and earnings conservatism and hence are unlikely to be driven by correlated omitted variables, measurement error or anomalies in any particular proxy.

V. THE ROLE OF ADDITIONAL INSTITUTIONAL FEATURES AND CAPITAL MARKET STRUCTURE

In this section, we examine several institutional factors, which we expect to differentially affect private and public firms. Specifically, we introduce binary indicator variables into our base model (Table 5, Model 3) to analyze whether the relation between earnings management and tax alignment, accrual accounting rules, securities regulation, outside investor protection, and capital market structure in fact differs across private and public firms. Unless indicated otherwise, we create these indicator variables by splitting at the median value of the institutional factors (see Table 3).

Financial Accounting and Tax Alignment

To capture differences in the tax regimes across EU countries, we rely on a classification provided by Alford et al. (1993) and Hung (2001). The classification indicates countries with a high alignment of financial and tax accounting, for instance, because financial statements serve as the basis for taxation or because tax laws explicitly require that certain items are treated equally in both sets of accounts. The *TAX* variable takes on a value of 1 when financial and tax accounts are highly aligned, and 0 otherwise. For the three countries with missing tax information (Austria, Greece, and Portugal), we assume a tax status of 1.[35]

Column one in Table 6 presents the results introducing the *TAX* variable. The coefficient on tax alignment has the expected sign and is significant over and above listing status and enforcement quality. The interaction between tax alignment and listing status is negative, as expected, but not statistically significant. To construct a finer measure of the extent to which there are tax incentives, we multiply the tax alignment factor with the average corporate tax rate (*IMD World Competitiveness Yearbook 2000*). Based on this combined metric, we create a dichotomous variable splitting the sample into two groups (high tax alignment/tax rate countries versus low tax alignment/tax rate countries).[36] For the former group, tax considerations are likely to play a larger role resulting in less informative earnings. Consistent with this claim, we find a positive and highly significant main effect for the adjusted tax alignment variable (column 2), suggesting that firms in high alignment/

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[35] The rationale for these assignments is that Austria is close to the German system, which is classified as a 1, and Portugal and Greece are both French legal origin countries, which generally exhibit a relatively high tax alignment. If we exclude the three countries without explicit classification from the tax analysis, the results do not materially change.

[36] We use a tax rate of 34 percent as cut-off value to account for the fact that the three countries with low tax alignment automatically get assigned to the base group. If we split by the median (i.e., reclassify Austria), the results are very similar although the interaction term is weaker in statistical significance.

TABLE 6

Earnings Management and Reporting Incentives: The Incremental Role of Tax Alignment, Accrual Accounting Rules, Securities Regulation, Investor Protection, and Capital Market Structure

$$EM_{aggr,i} = \alpha_0 + \alpha_1 \text{Conditional Variable}_i + \alpha_2 \text{Conditional Variable}_i{*}PUBL_i + \alpha_3 PUBL_i + \alpha_4 LEGAL_i + \alpha_5 SIZE_i + \alpha_6 LEV_i + \alpha_7 GROWTH_i + \alpha_8 ROA_i + \alpha_9 CYCLE_i + \Sigma\alpha_j \text{Industry Controls}_i + \varepsilon_i$$

	Tax Alignment		Accrual Accounting	Securities Regulation and Investor Protection		Capital Market Structure	
Variables	*TAX*	*TAX*RATE*	*ACCRUAL*	*SECREG*	*ANTIDIR*	*SYSTEM*	*DEV*
n	269	269	269	269	269	269	269
Conditional Variable	4.180#	16.554**	−7.354**	−0.306	−0.011	1.735	2.110
	(1.67)	(6.22)	(−3.08)	(−0.14)	(0.00)	(0.82)	(1.03)
Conditional Variable*PUBL*	−3.384	−8.772**	−1.844	−5.614	−9.372**	−6.931#	−8.511*
	(−0.94)	(−2.58)	(−0.50)	(−1.60)	(−2.61)	(−1.80)	(−2.44)
PUBL	−17.269**	−12.978**	−18.995**	−17.254**	−18.269**	−16.151**	−15.438**
	(−3.75)	(−3.06)	(−4.77)	(−3.80)	(−4.51)	(−3.78)	(−3.65)
LEGAL	−5.707**	−2.379#	−5.185**	−6.027**	−6.987**	−6.724**	−6.437**
	(−4.30)	(−1.67)	(−4.33)	(−5.09)	(−5.71)	(−5.48)	(−5.48)
SIZE	1.767	1.072	2.130	2.031	1.935	1.951	1.977
	(1.26)	(0.84)	(1.55)	(1.45)	(1.44)	(1.40)	(1.43)
LEV	27.787**	22.554**	20.907*	25.542**	29.793**	33.283**	29.351**
	(3.03)	(2.69)	(2.39)	(2.61)	(3.24)	(3.59)	(3.37)
GROWTH	−65.498*	−56.209*	−57.577*	−61.065*	−61.492*	−67.984*	−61.227*
	(−2.35)	(−1.97)	(−2.36)	(−2.34)	(−2.38)	(−2.48)	(−2.36)
ROA	−2.546**	−2.115**	−2.431**	−2.726**	−2.447**	−2.536**	−2.691**
	(−3.77)	(−3.27)	(−4.12)	(−4.13)	(−3.89)	(−3.94)	(−4.22)
CYCLE	0.079#	0.076#	0.053	0.073#	0.057	0.063	0.061
	(1.84)	(1.78)	(1.27)	(1.75)	(1.35)	(1.47)	(1.44)
Intercept and Industry Controls	included	included	included	included	included	included	included
R^2	59.9%	64.1%	62.3%	60.4%	60.8%	60.2%	60.7%

(continued on next page)

TABLE 6 (Continued)

**, *, and # Indicate statistical significance at the 1 percent, 5 percent, and 10 percent levels (two-tailed), respectively.

The sample comprises 269 industry-level observations from 13 European countries. The dependent variable, EM_{aggr}, is the average percentage rank across all four individual earnings management scores, *EM1* to *EM4*, as described in Table 1. *PUBL* is an indicator variable taking on the value of 1 if the observation stems from publicly traded firms, and 0 otherwise. *LEGAL* stands for the mean of three variables in La Porta et al. (1998), which measure the quality of the legal system and enforcement (i.e., efficiency of the judicial system, rule of law, and corruption index). The table reports results for the following conditional variables: *TAX* is an indicator variable taking on the value of 1 if financial accounts for external reporting and tax purposes are highly aligned, and 0 otherwise (see Alford et al. 1993; Hung 2001). We assume a tax status of 1 for the three countries with missing tax information (Austria, Greece, and Portugal). *TAX*RATE* is the tax alignment indicator multiplied by the average corporate tax rate in percent of earnings before taxes (Source: *IMD World Competitiveness Yearbook 2000*) thereby giving more weight to high tax rate countries. *ACCRUAL* is the accrual index from Hung (2001) (updated for European countries by Comprix et al. [2003]), and captures differences in accrual accounting rules across countries. *SECREG* captures the strength of securities regulation mandating and enforcing disclosures for publicly listed firms. It is measured as the mean of the disclosure index, the liability standard index, and the public enforcement index from La Porta et al. (2006). *ANTIDIR* is the antidirector rights index from La Porta et al. (1998) capturing the legal protection of minority shareholders. The capital market variables consist of two measures from Beck and Levine (2002): *SYSTEM* captures whether a country's financial system is more market- or bank-based, and equals the natural log of (total value of stock traded/claims on the private sector by commercial banks), both scaled by GDP. Financial development, *DEV*, is a measure of the overall activity of financial intermediaries and of capital market development. It equals the natural log of (total value of stock traded*claims on the private sector by financial institutions), both scaled by GDP. We transform the continuous institutional factors into binary variables splitting by the median (except for *TAX*RATE* where we use 34 percent as a cut-off, and *ANTIDIR* where we split by the value of 4; see discussion in the text for details). The model includes the main effects and the interaction term of the conditioning variable and listing status, the legal enforcement variable as well as the controls from the base specification (see Model 3 in Table 5). The table reports OLS coefficient estimates and t-statistics based on heteroscedasticity corrected standard errors (in parentheses).

tax rate countries engage in more earnings management. However, we also find a significantly negative interaction term, consistent with our hypothesis in Section II that capital market forces provide a counterweight to the influence of tax incentives.[37]

Remaining Differences in Accrual Accounting Rules

To capture remaining accounting differences in the EU, we use the "accrual rules index," *ACCRUAL*, constructed by Hung (2001) and updated for EU countries by Comprix et al. (2003). This index measures the use of accrual rules to accelerate recognition of economic transactions (e.g., R&D activities, pension plans), which is generally viewed as improving the informativeness of earnings (e.g., Dechow 1994). Higher index values correspond to more extensive accrual rules (see Table 3).

Column three in Table 6 examines the relation between a binary indicator for accrual accounting regimes and earnings management. The results show that *ACCRUAL* is significantly negative at the .01 level, and so are the two main incentives variables. Thus, our previous results are not adversely affected when we explicitly control for residual differences in EU accounting rules. As discussed in Section II, the effects of accrual accounting rules may differ across private and public firms. Rules that are designed to produce timely and informative earnings are likely to be more important for public firms, which rely heavily on financial statements to communicate performance to outside investors. The negative coefficient on the interaction term between *ACCRUAL* and *PUBL* points in this direction, but it is insignificant. Thus, standards that are intended to produce timely and informative earnings play an incremental role and appear to matter equally to all firms in an economy.[38]

Securities Regulation and Minority-Shareholder Protection

To capture differences in securities regulation, we create a variable using three indices from a recent study by La Porta et al. (2006): (1) the disclosure requirements index representing several aspects of prospectus disclosure in public security offerings, (2) the liability standard index measuring the procedural difficulties in recovering losses from the issuer and its directors in a civil liability case, and (3) the public enforcement index capturing market supervision by a regulator and its investigative powers and sanctions. By calculating the arithmetic mean of the three indices, we create a summary measure *SECREG* representing the overall effectiveness of a country's securities regulation (see Table 3). To capture minority-shareholder protection, we use the anti-director rights index from La Porta et al. (1998). This index aggregates several regulatory measures protecting shareholders' voting and supervisory rights and ranges from 0 to 6. Values of 4 or higher are generally viewed as representing high outside investor protection, which is why we use this cut-off to form our binary indicator variable (e.g., Leuz et al. 2005).

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[37] In further robustness tests (not tabulated) we re-run separate analyses either using observations from consolidated financial statement data or observations from unconsolidated (or parent-only) accounts because the alignment of tax and financial accounting is commonly based on the parent-only accounts. Even though we find a much larger and more significant positive main effect using the subsample of unconsolidated accounts, we do not find evidence of a differential tax effect between public and private firms for either of the two subsamples, which might be due to the substantial decrease in sample size. Moreover, when we re-run the analyses weighting *TAX* by a tax-compliance measure (using a survey metric of the extent of tax compliance in each country from the *2000 IMD World Competitiveness Yearbook*) instead of *RATE*, the results are similar, though the main effect of tax compliance is not significant.

[38] We further examine the above findings by either using consolidated or parent-only data. Untabulated analyses show that the main effect of the accrual rules variable for the consolidated sample is significant at the .02 level, consistent with the idea that consolidated numbers serve as the primary tool to convey information to company outsiders. Using parent-only accounts, on the other hand, the *ACCRUAL* variables are not statistically significant.

In Table 6, we present results introducing these two variables into our model. Column four shows that the interaction term between *SECREG* and *PUBL* is negative and marginally significant at the .11 level (two-sided) and that the main effect is insignificant. The latter result is expected, given that securities regulation is primarily geared toward public firms. The marginally significant interaction term suggests that strong securities regulation may reinforce the effect of capital markets in reducing earnings management. We obtain a similar but much stronger result when we use *ANTIDIR* (column 5). The interaction effect is now highly significant while the main effect remains insignificant, which is expected given that strong minority-shareholder rights matter primarily for arm's-length financing in public equity markets. We interpret these findings as evidence that capital market forces faced by public firms and legal institutions geared toward public firms reinforce each other in providing incentives to report earnings that reflect firms' economic performance.

Capital Market Structure

To further explore the effects of market forces, we examine two features that capture countries' capital market structure and financial development. The basic idea is that capital market incentives should be stronger in countries that already have active and highly developed public capital markets (see also, Bushman and Piotroski 2006). We employ two variables from Beck and Levine (2002) to capture these aspects and interact them with *PUBL*. *SYSTEM* measures whether the financial system is relatively more equity-market-based versus bank-based and, hence, the extent to which there is arm's length or relationship financing. It is computed as the natural log of the value of stock transactions in an economy divided by commercial banks' claims on the private sector. *DEV* is a measure of financial development and the overall capital market activity. It equals the natural log of the value of stock transactions in an economy times financial institutions' claims on the private sector (see Table 3).[39]

The last two columns in Table 6 present results introducing binary indicator variables for both capital market features into the model. The results are very similar across the two variables. In both cases, the interaction terms between the capital market features and *PUBL* are significantly negative, indicating that there is less earnings management among public firms in countries with equity-based and highly developed capital markets. Again, we find that these effects apply only to public firms, as might be expected.

Taken together, our analyses suggest that the first-order effect of public equity markets is to reduce earnings management.[40] This effect interacts with several institutional features in interesting ways. Capital market forces can mitigate the negative effects of tax alignment and reinforce the positive effects of legal institutions geared primarily toward public firms. Moreover, the strength of equity market forces depends on the structure and development of countries' capital markets in place.

VI. CONCLUSIONS

This study provides compelling evidence that firms' reporting incentives created by market forces and institutional factors are important determinants of accounting quality. We exploit the EU setting and a relatively new database to examine whether public equity markets and strong legal enforcement exert a systematic influence on firms' incentives to report earnings that accurately reflect economic performance.

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[39] If we use Wurgler's (2000) summary statistic of financial development instead, we obtain exactly the same country partition.

[40] In separate tests (untabulated), we repeat the analyses in Table 6 using our aggregate earnings informativeness index, EI_{aggr}, as the dependent variable (see Table 5, Panel B). With the exception of the interaction term in the *TAX*RATE* specification, all results are very similar and none of the inferences change.

We focus primarily on earnings management as one dimension of accounting quality and find for a large sample of private and public firms from 13 European countries that earnings management is more pervasive in privately held firms. These results are contrary to recent allegations that capital markets exacerbate incentives to mask economic performance. Instead, our findings suggest that the first-order effect of capital markets is to increase earnings informativeness. Capital markets may do so in several ways. For instance, the going public process may *ex ante* screen out firms with earnings that are innately less informative or difficult to evaluate by outsiders. In addition, capital markets are likely to provide incentives *ex post* to make earnings more informative. Our study does not attempt to identify the exact mechanism by which capital markets appear to curb earnings management. Given the substantial number of controls for firm characteristics in our analyses, we believe that *ex post* incentives are more likely to explain our findings than *ex ante* screening based on innate earnings quality, but leave this issue to future research.

We also provide evidence that earnings management is more pronounced in countries with weaker legal systems and enforcement. This effect is present for both public and private firms, confirming the central role of enforcement mechanisms. In addition, we explore the interaction between market forces and other institutional variables. The results generally support our primary finding that capital market forces improve the informativeness of earnings. We find that stronger tax alignment is associated with more earnings management, but that this effect is mitigated by market pressure on public firms. We find that legal institutions that facilitate equity financing in public markets, such as strong minority-shareholder rights and extensive disclosure requirements, reduce the level of earnings management primarily for publicly traded firms, suggesting that markets and institutions reinforce each other. The evidence also shows that in countries with large and highly developed equity markets, public firms engage in even less earnings management, providing further support for the notion that strong capital markets and arm's-length financing improve earnings informativeness.

While our results are robust to several alternative definitions of the dependent variable, to the inclusion of many different control and institutional variables, and to substantial modifications of the sample composition, we acknowledge that earnings management and earnings informativeness are notoriously difficult to measure. Thus, it is hard to preclude the possibility that our results are influenced by omitted variables, which is why we encourage future research to study the important issue of firms' reporting incentives in a variety of ways and settings.

REFERENCES

Alford, A., J. Jones, R. Leftwich, and M. Zmijewski. 1993. The relative informativeness of accounting disclosures in different countries. *Journal of Accounting Research* 31: 183–223.

Ali, A., and L. Hwang. 2000. Country-specific factors related to financial reporting and the value relevance of accounting data. *Journal of Accounting Research* 38: 1–21.

Ball, R., S. Kothari, and A. Robin. 2000. The effect of international institutional factors on properties of accounting earnings. *Journal of Accounting and Economics* 29: 1–51.

———. 2001. Infrastructure requirements for an economically efficient system of public financial reporting and disclosure. In *Brookings-Wharton Papers on Financial Services*, edited by R. Litan, and R. Herring. Washington, D.C.: Brookings Institution Press.

———, A. Robin, and J. Wu. 2003. Incentives versus standards: Properties of accounting income in four East Asian countries. *Journal of Accounting and Economics* 36: 235–270.

———, and L. Shivakumar. 2005. Earnings quality in U.K. private firms: Comparative loss recognition timeliness. *Journal of Accounting and Economics* 39: 83–128.

———, A. Robin, and G. Sadka. 2005. Is accounting conservatism due to debt or equity markets? An international test of "contracting" and "value relevance" theories of accounting. Working paper, University of Chicago.

Basu, S. 1997. The conservatism principle and the asymmetric timeliness of earnings. *Journal of Accounting and Economics* 24: 3–37.

Beatty, A., and D. Harris. 1999. The effects of taxes, agency costs and information asymmetry on earnings management: A comparison of public and private firms. *Review of Accounting Studies* 4: 299–326.

———, B. Ke, and K. Petroni. 2002. Earnings management to avoid earnings declines across publicly and privately held banks. *The Accounting Review* 77: 547–570.

Beck, T., and R. Levine. 2002. Industry growth and capital allocation: Does having a market- or bank-based system matter? *Journal of Financial Economics* 64: 147–180.

Bradshaw, M., and G. Miller. 2005. Will harmonizing accounting standards really harmonize accounting? Working paper, Harvard Business School.

Burgstahler, D., and I. Dichev. 1997. Earnings management to avoid earnings decreases and losses. *Journal of Accounting and Economics* 24: 99–126.

Bushman, R., J. Piotroski, and A. Smith. 2004. What determines corporate transparency? *Journal of Accounting Research* 42: 207–252.

———, and ———. 2006. Financial reporting incentives for conservative accounting: The influence of legal and political institutions. *Journal of Accounting and Economics* 42: 107–148.

Campbell, J. 1996. Understanding risk and return. *Journal of Political Economy* 104: 298–345.

Comprix, J., K. Muller III, and M. Stanford. 2003. Economic consequences from mandatory adoption of IASB standards in the European Union. Working paper, Arizona State University.

Coppens, L., and E. Peek. 2005. An analysis of earnings management by European private firms. *Journal of International Accounting, Auditing and Taxation* 14: 1–17.

Dechow, P. 1994. Accounting earnings and cash flows as measures of firm performance: The role of accounting accruals. *Journal of Accounting and Economics* 18: 3–42.

———, R. Sloan, and A. Sweeney. 1995. Detecting earnings management. *The Accounting Review* 70: 193–225.

———, and D. Skinner. 2000. Earnings management: Reconciling the views of accounting academics, practitioners, and regulators. *Accounting Horizons* 14: 235–250.

———, and C. Schrand. 2004. *Earnings Quality*. Charlottesville, VA: The Research Foundation of CFA Institute.

Degeorge, F., J. Patel, and R. Zeckhauser. 1999. Earnings management to exceed thresholds. *Journal of Business* 72: 1–33.

Fan, J., and T. Wong. 2002. Corporate ownership structure and the informativeness of accounting earnings in East Asia. *Journal of Accounting and Economics* 33: 401–425.

Fischer, P., and P. Stocken. 2004. Effect of investor speculation on earnings management. *Journal of Accounting Research* 42: 843–870.

Gernon, H., and R. Wallace. 1995. International accounting research: A review of its ecology, contending theories and methodologies. *Journal of Accounting Literature* 14: 54–106.

Haw, I., B. Hu, L. Hwang, and W. Wu. 2004. Ultimate ownership, income management, and legal and extra-legal institutions. *Journal of Accounting Research* 42: 423–462.

Healy, P., and J. Wahlen. 1999. A review of the earnings management literature and its implications for standard setting. *Accounting Horizons* 13: 365–383.

Hung, M. 2001. Accounting standards and value relevance of financial statements: An international analysis. *Journal of Accounting and Economics* 30: 401–420.

Jones, J. 1991. Earnings management during import relief investigations. *Journal of Accounting Research* 29: 193–228.

Joos, P., and M. Lang. 1994. The effects of accounting diversity: Evidence from the European Union. *Journal of Accounting Research* 32: 141–168.

———, and P. Wysocki. 2004. (Non)convergence in international accrual accounting: The role of institutional and real operating effects. Working paper, Sloan School of Management, MIT.

La Porta, R., F. Lopez-de-Silanes, A. Shleifer, and R. Vishny. 1998. Law and finance. *Journal of Political Economy* 106: 1113–1155.

———, ———, and ———. 1999. Corporate ownership around the world. *Journal of Finance* 54: 471–517.

———, ———, and ———. 2006. What works in securities laws? *Journal of Finance* 61: 1–32.

Land, J., and M. Lang. 2002. Empirical evidence on the evolution of international earnings. *The Accounting Review* 77: 115–133.

Lang, M., J. Smith Raedy, and M. Higgins Yetman. 2003. How representative are firms that are cross-listed in the United States? An analysis of accounting quality. *Journal of Accounting Research* 41: 363–386.

———, ———, and W. Wilson. 2006. Earnings management and cross-listing: Are reconciled earnings comparable to U.S. earnings? *Journal of Accounting and Economics* 42: 149–165.

Leuz, C., D. Nanda, and P. Wysocki. 2003. Earnings management and investor protection: An international comparison. *Journal of Financial Economics* 69: 505–527.

———, K. Lins, and F. Warnock. 2005. Do foreigners invest less in poorly governed firms? Working paper, European Corporate Governance Institute.

Levitt, A., Jr. 1998. The importance of high-quality accounting standards. *Accounting Horizons* 12: 79–82.

Myers, J., L. Myers, and D. Skinner. 2006. Earnings momentum and earnings management. Working paper, University of Chicago.

Peek, E., R. Cuijpers, and W. Buijink. 2006. Creditors' and shareholders' demand for accounting conservatism in public versus private firms: Evidence from Europe. Working paper, Maastricht University.

Rajan, R., and L. Zingales. 1995. What do we know about capital structure? Some evidence from international data. *Journal of Finance* 50: 1421–1460.

Roychowdhury, S. 2005. Management of earnings through the manipulation of real activities that affect cash flow from operations. Working paper, Sloan School of Management, MIT.

Saudagaran, S., and G. Meek. 1997. A review of research on the relationship between international capital markets and financial reporting by multinational firms. *Journal of Accounting Literature* 16: 127–159.

Sloan, R. 1996. Do stock prices fully reflect information in accruals and cash flows about future earnings? *The Accounting Review* 71: 289–315.

Stolowy, H., and A. Jeny-Cazavan. 2001. International accounting disharmony: The case of intangibles. *Accounting, Auditing and Accountability* 14: 477–496.

Subramanyam, K. 1996. The pricing of discretionary accruals. *Journal of Accounting and Economics* 22: 249–281.

Teoh, S., I. Welch, and T. Wong. 1998a. Earnings management and the underperformance of seasoned equity offerings. *Journal of Financial Economics* 50: 63–99.

———, ———, and ———. 1998b. Earnings management and the long-run market performance of initial public offerings. *Journal of Finance* 53: 1935–1974.

Titman, S., and R. Wessels. 1988. The determinants of capital structure choice. *Journal of Finance* 43: 1–19.

Tweedie, D., and T. Seidenstein. 2005. Setting a global standard: The case for accounting convergence. *Northwestern Journal of International Law and Business* 25: 589–608.

Van Hulle, K. 2004. From accounting directives to international accounting standards. In *The Economics and Politics of Accounting*, edited by C. Leuz, D. Pfaff, and A. Hopwood. Oxford, U.K.: Oxford University Press.

Vander Bauwhede, H., M. Willekens, and A. Gaeremynck. 2003. Audit firm size, public ownership, and firms' discretionary accruals management. *International Journal of Accounting* 38: 1–22.

Watts, R., and J. Zimmerman. 1986. *Positive Accounting Theory*. Englewood Cliffs, NJ: Prentice Hall.

Wurgler, J. 2000. Financial markets and the allocation of capital. *Journal of Financial Economics* 58: 187–214.

Wysocki, P. 2004. Discussion of ultimate ownership, income management, and legal and extra-legal institutions. *Journal of Accounting Research* 42: 463–474.

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The Information Content of SEC Filings and Information Environment: A Variance Decomposition Analysis

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ABSTRACT: Using the Vuolteenaho (2002) variance decomposition methodology, this study assesses the relative value relevance of cash flow, accrual, and expected return news on SEC and preliminary earnings filing dates, as measured by their contribution to the volatility of *unexpected* returns. Cash flow news is found to be more value-relevant than accrual news. Although expected return (risk) news is the least value-relevant, it is significantly correlated with *changes* in betas and returns at the preliminary and SEC filing dates, indicating association with changes in firm risk and discount rates. This study also documents that these informational components contain less (more) value-relevant information at the SEC filing date for firms with a higher proportion of long-term (short-term) sophisticated investors after controlling for other dimensions of the information environment.

Keywords: *SEC filings; value relevance; variance decomposition.*

I. INTRODUCTION

A number of papers have examined the value relevance of SEC filings for firms that provide preliminary earnings. Although most early studies tend to find little if any value relevance on SEC filing dates, possibly because of sample size limitations and the difficulty in pinpointing the timing of SEC information prior to EDGAR, more

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recent studies conclude that SEC filings are in fact value-relevant beyond preliminary earnings.[1] The twofold purpose of this study is to assess the relative value relevance of specific informational components on SEC filing (and preliminary earnings) dates and to investigate the extent to which these components are affected by the firm's information environment. Three informational components are of particular interest: news about the firm's risk as reflected in expected future returns (discount rates), news about the firm's accruals, and news about the firm's cash flows.

In contrast to prior studies, value relevance of a news component in this study is measured by its contribution to the *volatility* of *unexpected* equity returns. Information that is value-relevant necessarily causes market participants to revise expectations regarding discount rates (expected return news) and/or future cash flows (earnings news), thereby driving the volatility of current period returns.[2] This study initially focuses on the extent to which the volatility of unexpected equity returns around the SEC filing date is related to news concerning these two factors. We next decompose earnings news into accrual and cash flow news components to determine which of expected return (risk), accruals, or cash flow news most affect the volatility of unexpected equity returns around the SEC filing date. This decomposition is of special interest since accrual/cash flow news components are usually unknown at the time of the preliminary earnings announcement, and revealed for the first time in the SEC filings, in contrast to earnings and expected return (risk) news that are partially revealed at the preliminary earnings announcement.[3]

We assess the relative value relevance of the informational components at the SEC filing date (and at the preliminary date) using the variance decomposition methodology of Campbell (1991), Vuolteenaho (2002), and Callen and Segal (2004). The primary usefulness of these models is that they allow discount rates to change over time and, therefore, yield an estimable measure of the informational impact of SEC filings on shocks to firm risk as reflected in revisions to expected future discount rates. This is important because as Campbell et al. (1997, 265), among others, show, small changes in expected discount rates can have a large impact on security returns, especially when expected returns are persistent.[4] Of almost equal importance is that these models yield news estimates that are comprehensive. For example, earnings news in this study is measured as the impact of informational shocks on current period expected earnings (the earnings surprise) *and* on all expected future earnings over the lifetime of the firm. So for example, if the information in the SEC filing affects expectations of both current and future years' earnings, then earnings news will be the (discounted) sum of the current year earnings surprise and the change in expectations of future period's earnings.[5]

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[1] Foster and Vickrey (1978), Wilson (1987), Qi et al. (2000), Griffin (2003), and Balsam et al. (2002) document that 10-Q/Ks contain information beyond preliminary earnings announcements. Foster et al. (1983, 1986), Cready and Mynatt (1991), Stice (1991), Easton and Zmijewski (1993), and Chung et al. (2003) provide evidence that stock returns around SEC filing dates are not different from other days.

[2] Studies that have emphasized the volatility of unexpected equity returns as a valuation metric include Campbell (1991), Vuolteenaho (2002), Callen and Segal (2004), and Callen et al. (2005). See also Campbell et al. (1997) and Cochrane (2001).

[3] A few firms provide cash flow and/or accruals information in their preliminary earnings release. We control for the level of information on the preliminary earnings release date in the empirical tests below.

[4] We show below that expected return news is positively correlated with changes in firm risk and discount rates as measured by market betas and *ex post* returns.

The second major (interrelated) focus of this study concerns the relation between the informational components of unexpected equity return volatility on SEC filing dates and the informational environment of the firm, with a special focus on the degree of information asymmetry between the firm and its investors. It has been argued that disclosure regulation is motivated by regulators' desire to reduce the information gap between the firm and its investors and between informed and uninformed investors (Healy and Palepu 2001). If so, then a positive relation between the valuation relevance of SEC filings and the degree of information asymmetry should obtain. We measure information asymmetry in part by the degree of investor sophistication after controlling for other dimensions of the information environment, assuming that sophisticated investors are more knowledgeable than unsophisticated investors about the firm prior to SEC filings. An immediate consequence of this assumption is that sophisticated investors are less likely to revise their future earnings estimates based on the SEC filings as compared to less sophisticated investors, so that the contributions of expected return (risk), accrual, and cash flow news to the variance of unexpected returns on SEC filing dates should be negatively associated with the degree of investor sophistication.

We further break down sophisticated investors into two classes, investors with short-term investment horizons and investors with long-term investment horizons, following Bushee (1998, 2001). Not all sophisticated investors are likely to be equally motivated to collect information between the preliminary earnings announcement and the filing date. In particular, we expect the negative association between the variance contributions of expected return (risk) news, accrual news, and cash flow news on SEC filing dates and the degree of investor sophistication to hold primarily for sophisticated investors with long-term investment goals. No such relationship is expected for short-term ("transient") sophisticated investors, such as momentum traders, who have little incentive to collect additional firm-level information. On the contrary, momentum traders are more likely to view excess volatility in a positive light, since the extant empirical evidence suggests that institutional momentum traders invest primarily in smaller riskier firms (Bennett et al. 2000; Badrinath and Wahal 2002).

We find that the variance contribution of the expected return (risk) news component and the earnings news component are statistically significantly different from zero both at the preliminary earnings release date *and* the SEC filing date. These news items are also significantly negatively correlated. Despite the fact that earnings news is more economically significant than expected return (risk) news in driving unexpected return volatility at both the preliminary and SEC filing dates, nevertheless, expected return (risk) news is shown to be significantly associated with changes in betas and changes in *ex post* returns at the preliminary and SEC filing dates. Therefore, expected return news measures changes in firm risk and signals the direction of change in the firm's cost of capital at both the preliminary announcement and SEC filing dates.

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[5] Recent papers recognize the importance of controlling for the impact of current shocks on future earnings by controlling for persistency in the analysis, Yet, the earnings news definition and the VAR matrix in our analysis incorporates, by construction, the persistence of returns, the persistence of earnings (accruals/cash flows), and their interactions. We also do not have to control for firm growth, or interactions between firm growth and the various persistence measures because firm growth is incorporated in our VAR matrix (through the book-to-market ratio). For a proof of these claims in a simple environment, see Callen et al. (2005, Appendix A). For more general results, see Callen (2004).

We also document that the variance contribution of expected return (risk) news, accruals news, and cash flow news around the SEC filing date are statistically significantly different from zero, indicating that the significant variance contribution of earnings news on the SEC filing date may be attributed (at least in part) to the information conveyed by the accruals-cash flow breakdown. We also find that cash flow news is statistically and economically more significant than accrual news in driving volatility of returns. This finding contrasts with Callen and Segal (2004) who find that accruals and cash flows are equally important in driving unexpected *annual* returns.

This study further documents a negative association between the variance contribution of cash flow news and accruals news (but not expected return news) on the SEC filing date and the degree of investor sophistication as measured by aggregate institutional ownership. Further disaggregating sophisticated investors into long-term (dedicated and quasi-indexer) investors and short-term (transient) investors, reveals that expected return (risk) news, cash flow news, and accruals news on the SEC filing date are negatively correlated with the degree of investor sophistication for long-term investors. In contrast, we find a significantly positive association between the variance contribution of expected return (risk) and cash flows news components on the SEC filing date and the degree of investor sophistication for "transient" investors with short-term horizons.

In what follows, Section II develops the research hypotheses. Section III describes the variance decomposition model used to estimate earnings, accrual, and cash flow news. Section IV discusses the data and sample selection. Section V presents the empirical results and Section VI concludes.

II. HYPOTHESES DEVELOPMENT

Volatility Reactions to Preliminary Earnings Announcements and SEC Filing Dates

Prior research shows that investors respond to earnings surprises included in preliminary earnings announcements and that this information is incorporated fairly quickly into equity prices (Lev 1989; Kothari 2001). The accounting literature also shows that markets respond to SEC filings (Foster and Vickrey 1978; Wilson 1986, 1987; Qi et al. 2000; Griffin 2003; Balsam et al. 2002). In contrast to the prior literature, this study measures value relevance in terms of the volatility of unexpected equity returns. Of particular interest is the decomposition of volatility into informational components on the preliminary earnings announcement and SEC filing dates. Initially, two components are of interest, namely, earnings news and expected return (risk) news as formalized in the following two hypotheses, stated in the alternative:

H1: The contributions of expected return (risk) news and earnings news to the volatility of unexpected equity returns are significantly different from zero on the preliminary earnings announcement date.

H2: The contributions of expected return (risk) news and earnings news to the volatility of unexpected equity returns are significantly different from zero on the SEC filing date.

Expected Return (Risk), Operating Cash Flow, and Accruals News

When a company releases preliminary earnings prior to the SEC filing, the earnings news will be impounded in price. Therefore, one should not observe earnings news on the SEC filing date, unless the filed earnings are materially different from preliminary earnings or unless the information contained in the financial report causes investors to interpret the earnings number differently. One possible reason for the reinterpretation of the earnings number is the additional information conveyed by the breakdown of earnings into cash flows and accruals.

The literature on cash flows and accruals indicates that both components of earnings are generally value-relevant (Rayburn 1986; Wilson 1986, 1987; Bowen et al. 1987; Bernard and Stober 1989; Livnat and Zarowin 1990; Dechow 1994; Sloan 1996; Subramanyam 1996; Pfeiffer and Elgers 1999; Barth et al. 1999; Xie 2001; Callen and Segal 2004). Which of the two earnings components is more value-relevant remains unclear (Sloan 1999). Particularly germane to our study are the papers by Wilson (1986, 1987) and Callen and Segal (2004). Wilson (1986) examines market reactions to new information about cash flows from operations (or funds from operations) and accruals around the filing of the ARS or Form 10-K with the SEC. He finds significant market reactions to cash flows and accruals together for a small sample of manufacturing firms. He also finds that accruals are value-relevant beyond cash flows. Wilson (1986) does not evaluate the impact of the accruals-cash flow breakdown on the volatility of unexpected equity returns nor does he measure changes in firm risk. Callen and Segal (2004) employ the variance decomposition methodology to examine the variance contribution of expected return (risk), accruals, and cash flows news on the volatility of unexpected *annual* returns. They find that accruals and cash flow news are far more value-relevant than expected return (risk) news. They also find some evidence that accrual news is more value-relevant than cash flow news. However, they do not investigate the return decomposition of expected return (risk), accruals, and net operating cash flow news over the short window around the SEC filing date, when this information initially becomes available to investors.

This discussion leads to our third hypothesis (stated in the alternative):

H3: The contributions of expected return (risk), accruals, and cash flows news to the volatility of unexpected equity returns on the SEC filing date are significantly different from zero.

Earnings News and Investor Sophistication

The greater the degree of information asymmetry between the firm and its investors, the more likely are preliminary earnings and SEC filings to be informative to the equity markets. Recent work by Balsam et al. (2002) and Griffin (2003) point to a negative relation between the information content of the financial statements on SEC filing dates and investor sophistication. This suggests that the informational components of the volatility of unexpected equity returns at the SEC filing date are a decreasing function of investor sophistication, after controlling for other dimensions of the information environment. This is expressed formally in our next hypothesis (stated in the alternative):

H4: The variance contributions of expected return (risk) news, cash flow news, and accrual news on the SEC filing dates are decreasing with the level of investor sophistication.

Bushee (1998, 2001) classifies institutions into three groups—"transient," "dedicated," and "quasi-indexers"—based on factors such as portfolio turnover, diversification, and momentum trading. "Transient" institutions are characterized as having high portfolio turnover and highly diversified portfolio holdings. These are short-term focused investors whose investments are based on the likelihood of short-term trading profits and have little incentive to gather information relevant to long-run value. In contrast, "dedicated" investors and "quasi-indexers" both provide long-term, stable ownership to firms. Dedicated institutional investors hold large stakes in a limited number of firms, whereas quasi-indexers generally follow indexing and buy-and-hold strategies, characterized by high diversification. Such ownership structures create incentives to invest in monitoring management and to obtain information beyond current earnings to assess managers' performance.[6] In addition, Porter (1992) argues that long-term investors have better access to private information about their portfolio firms. Thus, greater proportions of "dedicated" investors and "quasi-indexers" are associated with lower information asymmetry. Given the divergence in investor objectives and information needs, our final hypothesis (stated in the alternative) examines whether the level of ownership by different institutional investors has implications for the variance contributions of the news components.

> **H5:** The variance contributions of expected return (risk) news, cash flow news, and accruals news on the SEC filing dates are decreasing (not decreasing) with the level of investment by long-term (short-term) institutional investors.

III. NEWS, VARIANCE DECOMPOSITION, AND RESEARCH DESIGN

This section briefly describes the Vuolteenaho (2002) and Callen and Segal (2004) valuation models and the variance decomposition methodology. The method for estimating the news components is relegated to an Appendix.

The Models

Based on the Clean Surplus relation, Vuolteenaho (2002) shows that a security's unexpected returns can be expressed as:[7]

$$r_t - E_{t-1}(r_t) = \Delta E_t \sum_{j=0}^{\infty} \rho^j (roe_{t+j} - i_{t+j}) - \Delta E_t \sum_{j=1}^{\infty} \rho^j r_{t+j} \tag{1}$$

where:

Δ = the first differencing operator;
E_t = the expectations operator and $\Delta E_t = E_t(\cdot) - E_{t-1}(\cdot)$;
r_t = log equity return (cum dividend) in excess of the risk-free rate in period t;
ρ = a constant error approximation term;
roe_t = log book return on equity in period t = $\log(1 + X_t/BV_{t-1})$;
BV_t = book value of equity at time t;
X_t = net income in period t; and
i_t = log of 1 plus the risk-free rate in period t.

Equation (1) shows that the shock to returns $(r_t - E_{t-1}(r_{t-1}))$ depends not only on the shock to current period (log deflated) expected earnings—the conventional earnings surprise $\Delta E_t(roe_t)$—but also on the shock to future expected earnings $\Delta E_t \sum_{j=1}^{\infty} \rho^j (roe_{t+j})$. Defining the unexpected stock return components as expected-return (risk) news (*Nr*) and earnings news (*Ne*), Equation (1) can be expressed as:

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[6] Although quasi-indexers follow a passive investment strategy, Monks and Minow (1995) argue that these investors have strong incentives to monitor management to ensure that it is acting in the best interest of the firm.

[7] Small case letters denote the logs of the capitalized variables. For simplicity, we abstract from any additive error approximation terms when describing the models.

$$r_t - E_{t-1}(r_t) = Ne_t - Nr_t \tag{2}$$

where:

$$Nr_t = \Delta E_t \sum_{j=1}^{\infty} \rho^j r_{t+j} = \text{Expected Return News; and} \tag{3}$$

$$Ne_t = \Delta E_t \sum_{j=0}^{\infty} \rho^j (roe_{t+j} - i_{t+j}) = \text{Earnings News.} \tag{4}$$

Equation (2) shows that the unexpected change in current stock returns increases with earnings news and decreases with expected return news. An unexpected positive change in earnings news conveys positive information about the firm's prospects and, hence, translates into higher returns. Conversely, an unexpected increase in future expected returns (costs of capital), due to increased risk, translates into negative unexpected current returns.

Taking variances of both sides of Equation (2) yields the variance decomposition of unexpected stock returns:

$$\text{Var}[r_t - E_{t-1}(r_t)] = \text{var}(Nr_t) + \text{var}(Ne_t) - 2\text{cov}(Nr_t, Ne_t). \tag{5}$$

Equation (5) can be used to assess the impact of earnings news and expected-return (risk) news on equity returns. The greater is the variance of any factor on the right-hand side of the equation, the more impact that factor has in explaining the volatility of unexpected returns.

The impact of cash flow news and accrual news can be determined in similar fashion. Following Callen and Segal (2004), the log return on book equity in Equation (1) can be broken down into cash flow and accrual components yielding:[8]

$$r_t - E_{t-1}(r_t) = \Delta E_t \sum_{j=0}^{\infty} \rho^j (CF_{t+j} - i_{t+j}) + \Delta E_t \sum_{j=0}^{\infty} \rho^j ACC_{t+j} - \Delta E_t \sum_{j=1}^{\infty} \rho^j r_{t+j} \tag{6}$$

where CF_t denotes cash flow from operations normalized by prior period book value and ACC_t denotes accruals normalized by prior period book value. Equation (6) can be expressed more succinctly as:

$$r_t - E_{t-1}(r_t) = nCF_t + nACC_t - Nr_t \tag{7}$$

where:

$$nCF_t = \Delta E_t \sum_{j=0}^{\infty} \rho^j (CF_{t+j} - i_{t+j}) \text{ Cash Flow News; and} \tag{8}$$

$$nACC_t = \Delta E_t \sum_{j=0}^{\infty} \rho^j ACC_{t+j} = \text{Accrual News.} \tag{9}$$

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[8] It is immaterial whether i_t is subtracted from cash flows or accruals in Equation (6).

Equation (7) can be used to provide a variance decomposition of the unexpected change in returns based on the cash flow-accrual breakdown. Specifically, taking variances of both sides of this equation yields:

$$\begin{aligned}\text{var}[r_t - E_{t-1}(r_t)] &= \text{var}(Nr_t) + \text{var}(nCF_t) + \text{var}(nACC_t) - 2\text{cov}(Nr_t, nCF_t)\\ &\quad - 2\text{cov}(Nr_t, nACC_t) + 2\text{cov}(nCF_t, nACC_t). \qquad (10)\end{aligned}$$

The VAR

The return variance decomposition (Equation (5)) cannot be computed without estimates (of the dynamics) of expected returns and expected earnings. Following Campbell (1991), Campbell and Ammer (1993), Vuolteenaho (2002), Callen (2004), Callen and Segal (2004), and Callen et al. (2005), we estimate earnings and expected returns news using a log-linear vector autoregressive (VAR) model. In general, VAR estimation is facilitated by assuming that the dynamics of the data are well described by a (stationary) multivariate time-series model. Specifically, define z_{it} to be a vector of firm-specific state variables that follows the vector autoregressive process:

$$z_{i,t} = Az_{i,t-1} + \eta_{i,t}. \qquad (11)$$

Consistent with the literature, the VAR coefficient matrix **A** is assumed to be constant over time and over firms. The error term vectors $\eta_{i,t}$ are zero-mean vectors of shocks and are assumed to have a variance-covariance matrix Ω and to be independent of all variables known at $t-1$.

We initially estimate a parsimonious VAR where the state variables consist of log stock returns, log book return on equity, and the log book-to-market ratio.[9] The VAR model can then be described as a system of (mean-adjusted) equations:

$$r_t = \alpha_1 r_{t-1} + \alpha_2 roe_{t-1} + \alpha_3 bm_{t-1} + \eta_{1t}; \qquad (12a)$$

$$roe_t = \beta_1 r_{t-1} + \beta_2 roe_{t-1} + \beta_3 bm_{t-1} + \eta_{2t}; \text{ and} \qquad (12b)$$

$$bm_t = \delta_1 r_{t-1} + \delta_2 roe_{t-1} + \delta_3 bm_{t-1} + \eta_{3t}. \qquad (12c)$$

We estimate the regressions separately using panel data with firm and fiscal quarter fixed effects. Consistent robust standard errors are obtained using the Shao and Rao (1993) jackknife method where the clustering variable is the calendar year fiscal quarter.

IV. DATA AND SAMPLE

The Preliminary and Unrestated Compustat Quarterly Data

Data entry into the Compustat databases has been performed in a fairly structured manner over the years. When a firm releases its preliminary earnings announcement, Compustat takes as many line items as possible from the preliminary announcement and enters them into the quarterly database within two to three days of the announcement. The preliminary data in the database are denoted by an update code of 2, until the firm files its Form 10-Q (10-K) with the SEC or releases it to the public, at which point Compustat updates all available information and uses an update code of 3. In subsequent periods,

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[9] As in Vuolteenaho (2002) the book-to-market ratio is included in the parsimonious VAR because our model is generated from this ratio. It also controls for the firm's growth prospects.

Compustat updates the information further if the firm restates its quarterly financial statements for any reason. Most restatements are due to mergers and acquisitions or divestitures.

Charter Oak Investment Systems, Inc. (Charter Oak) has routinely collected the weekly original CD-ROMs that Compustat sends to its PC clients, which contain updated data as of that week. From these weekly updates, Charter Oak constructs a database that contains, for each firm in the Compustat Quarterly database, three numbers for each Compustat line item in each quarter. The first number is the preliminary earnings announcement that Compustat inserts into the database bearing the update code of 2. The second number is the "As First Reported" (AFR) figure when Compustat first changes the update code to 3 for that firm-quarter. The third number is the number that exists in the current version of Compustat, which is what most investors and academics use. The Charter Oak database allows us to determine the information contained in the preliminary earnings announcement and the first-reported information in the SEC filing. Therefore, the Charter Oak database is superior to Compustat because the Charter Oak earnings number is what investors see and react to on the SEC filing date, untainted by restatements. Also, and unlike Compustat, the Charter Oak database provides data that allow us to control for those firms that disclose more than just earnings on the preliminary earnings announcement date.[10]

Sample Selection

The initial population for this study consists of all firm-quarters in the Quarterly Compustat database between the first quarter of 1988 (the first quarter after the adoption of SFAS No. 95, which mandated the disclosure of net operating cash flows) and the most recent quarter available at the time of the study (the fourth quarter of 2003). The only limitation that we impose on the initial selection of firm-quarters is to require that firm market values at quarter-end be in excess of $1 million, yielding an initial population of 348,275 firm quarters. From this initial population, we delete observations with missing CUSIP identifier or missing preliminary earnings release dates. In addition, observations are deleted if the preliminary income before extraordinary items and discontinued operations (DATA 8) is missing. We also delete observations with missing quarterly net operating cash flow in the SEC filing (DATA 108), missing total assets at the end of the prior quarter (DATA 44), and missing or negative book value of equity at the end of the prior quarter (DATA 60). These restrictions reduce the sample size to 325,525 firm-quarters.

For each firm-quarter in the reduced sample, we obtain the SEC filing date, which is supplied to us by Compustat for the calendar years 1991–2003. For each firm-quarter, we compute the cumulative return during the three-day window centered on the preliminary earnings announcement date and on the SEC filing date, as long as the two dates are in excess of seven days apart. The requirements of SEC filing dates (which are at least seven days after the preliminary announcement) and availability of returns around the preliminary earnings and SEC filing dates reduce the sample size further to 122,911 firm-quarter observations.

In addition, for each firm-quarter we compute the return on equity (ROE) on the preliminary (filing) date as preliminary (filing) income before extraordinary items and discontinued operations scaled by book value of equity at the end of the previous quarter. The book-to-market ratio (BM) is computed as book value of equity at the end of the previous quarter scaled by the contemporaneous market value of equity. Accruals are computed as income before extraordinary items and discontinued operations minus net operating cash flow.

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[10] We repeated the analysis using Compustat and obtained similar results.

To facilitate the estimation of the VAR earnings model on the preliminary earnings release date and the SEC filing date, as well as the VAR cash flow-accruals model on the SEC filing date, we require that each firm-quarter have both a valid current and lagged observation, where the lag refers to the same fiscal quarter in the previous fiscal year, for the following variables: cumulative returns, ROE and BM around the preliminary and filing date, and accruals and net operating cash flows around the filing date. These restrictions reduce the sample size to 85,254 firm-quarters. In addition, to mitigate the effect of outliers, we eliminate the top and bottom one percentile of each of the VAR system variables (current and lag), thereby reducing the sample size to 72,533 firm-quarters. Finally, we eliminate observations for which income at the preliminary date is materially different from the reported income in the 10-Q/-K because of their potential effect on earnings news at the filing date due to new information. Specifically, we drop observations for which the absolute value of the difference between income before extraordinary items at the filing date and preliminary date is greater than 1 percent of the absolute value of income before extraordinary items at the preliminary date, yielding a final sample of 70,606 firm-quarters.[11]

Sample Summary Data

Most firms disclose their preliminary earnings for the quarter or the year through a press release, followed by an SEC filing several weeks later. Easton and Zmijewski (1993) report a median lag between the fiscal quarter-end and the preliminary earnings announcement date (SEC filing date) of 28 (45) days. Our sample shows (untabulated) a similar pattern with median lags between the fiscal quarter-end and the preliminary earnings and SEC filing dates of 22 and 45 days, respectively.

Table 1 provides descriptive statistics of the main variables used in the analyses. The sample firms are large, with a mean (median) market value of equity of \$2.8 (\$0.3) billion. The table also reveals considerable variation and skewness in firm size. The sample firms have mean (median) quarterly return on equity of 2.1 percent (2.9 percent) and book-to-market ratio of 0.6 (0.5). The mean (median) accruals scaled by book value is −2.9 percent (−2.3 percent), and the mean and median net operating cash flows scaled by book value are 5.0 percent and 4.8 percent, respectively. The mean (median) three-day cumulative raw equity market return centered on the preliminary earnings release date of 0.8 percent (0.4 percent) is significantly larger than the three-day cumulative return centered on the 10-Q/-K filing date of 0.3 percent (0.1 percent). On average, about 29 percent of sample firms provide some balance sheet data at the preliminary earnings announcement but only about 8 percent provide cash flow data.

V. EMPIRICAL RESULTS

Variance Decomposition and Informational Components

The parameter estimates for the preliminary VAR earnings model are shown in Panel A of Table 2. The first number is the estimated coefficient based on firm and fiscal quarter fixed-effects panel data techniques. The second number in parentheses is a robust Shao and Rao (1993) jackknife standard error. All parameter estimates are significant at less than the

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[11] Inclusion of these observations in the analyses has no material impact on the reported results. A separate variance decomposition analysis for the sample of observations for which income at the preliminary date is materially different from the reported income in the 10-Q/-K shows a higher variance contribution of expected return news on the SEC filing date than that of firms with no earnings revisions between the preliminary earnings release and the SEC filing. These findings suggest that those firms that report a different earnings number in the SEC filing than in the preliminary earnings release experience an increase in risk as reflected by expected return news.

TABLE 1
Descriptive Statistics

	Mean	Std	Q1	Median	Q3
MV	2,804	14,693	95	332	1,223
R_P	0.008	0.071	−0.029	0.004	0.043
R_F	0.003	0.045	−0.021	0.001	0.025
BM	0.609	0.416	0.314	0.509	0.786
ROE	0.021	0.057	0.009	0.029	0.047
CF_B	0.050	0.095	0.006	0.048	0.092
AC_B	−0.029	0.090	−0.064	−0.023	0.011
EDGAR	0.710	0.454	0.000	1.000	1.000
FD	0.327	0.469	0.000	0.000	1.000
ANN	0.279	0.448	0.000	0.000	1.000
ACC_PROP	0.031	0.023	0.017	0.025	0.038
PREBS	0.287	0.452	0.000	0.000	1.000
PREOCF	0.077	0.267	0.000	0.000	0.000
LOSS	3.032	4.296	0.000	1.000	4.000
INS_HOL	0.471	0.239	0.282	0.482	0.662
DED_HOL	0.113	0.100	0.031	0.094	0.169
TRA_HOL	0.132	0.126	0.033	0.094	0.197
QIX_HOL	0.219	0.136	0.107	0.208	0.313

This table provides descriptive statistics on the main variables used in the analysis for our sample of 70,606 firm-quarters.

The institutional ownership statistics are based on a sample of 55,653 observations.

MV = market value of equity in millions of dollars at quarter-end;
R_P and R_F = the three-day cumulative raw returns centered on the preliminary earnings release date and on the 10-Q/K filing date, respectively;
BM = book value of equity at previous quarter-end divided by market value of equity at quarter-end;
ROE = return on equity, computed as the quarterly net income before extraordinary items divided by book value of equity at the previous quarter-end;
CF_B = quarterly cash flows from operations scaled by book value of equity at the previous quarter-end;
AC_B = accruals, computed as net income before extraordinary items minus quarterly cash flows from operations, scaled by book value of equity at the previous quarter-end;
EDGAR (FD) = dummy variable that equals 1 after 1995 (2000), and 0 otherwise;
ANN = dummy variable that equals 1 if a 10-K, and 0 otherwise;
ACC_PROP = the absolute value of the ratio of total accruals to average total assets in the prior eight quarters;
PREOCF (PREBS) = dummy variable that equals 1 if the company provided operating cash flows (balance sheet items on receivables, payables, and inventories) at the preliminary earning release date, and 0 otherwise;
LOSS = the number of loss quarters during the most recent 20 quarters;
INS_HOL = institutional ownership;
DED_HOL = the proportion of shares held by "dedicated" institutional investors; and
QIX_HOL (TRA_HOL) = proportion of shares held by "quasi-indexer" ("transient") institutional investors.

1 percent significance level. The return variable (r_t) is the three-day cumulative log equity return centered on the preliminary earnings announcement date. The return equation shows that current returns are negatively associated with past returns and past preliminary earnings (measured as the log of 1 plus the return on lagged book equity) and positively associated with past (log) book-to-market ratios. Thus, all three lagged variables Granger-cause returns.

TABLE 2
Estimated Parameters of the VAR Models

Panel A: Preliminary Earnings Model

VAR Coefficient Matrix: Preliminary Earnings

	r_{t-1}	roe_{t-1}	bm_{t-1}	Adj R^2	F-value
r_t	−0.041*** (0.006)	−0.053*** (0.009)	0.004*** (0.001)	0.003	56
roe_t	0.028*** (0.004)	0.205*** (0.016)	−0.006*** (0.001)	0.232	556
bm_t	−0.543*** (0.043)	−0.609*** (0.083)	0.395*** (0.025)	0.606	2471

Variance Decomposition: Preliminary Earnings

VARNRP	*VARNEP*	*COVRPE*	*VARRP*
0.000022*** (0.000001)	0.004076*** (0.000128)	−0.000146*** (0.000005)	0.004391*** (0.000138)

Panel B: Filing Date Earnings Model

VAR Coefficient Matrix: Filing Date Earnings

	r_{t-1}	roe_{t-1}	bm_{t-1}	Adj R^2	F-value
r_t	−0.042*** (0.008)	−0.014** (0.006)	0.001 (0.001)	0.000	33
roe_t	0.010* (0.006)	0.209*** (0.016)	−0.006*** (0.001)	0.234	538
bm_t	−0.340*** (0.078)	−0.688*** (0.086)	0.389*** (0.025)	0.601	2352

Variance Decomposition: Filing Date Earnings

VARNRF	*VARNEF*	*COVRFE*	*VARRF*
0.000004*** (0.000000)	0.001617*** (0.000055)	−0.000065*** (0.000002)	0.001751*** (0.000060)

Panel C: Cash Flow-Accrual Model

VAR Coefficient Matrix: Cash Flow-Accrual

	r_{t-1}	CF_{t-1}	ACC_{t-1}	bm_{t-1}	Adj R^2	F-value
r_t	−0.041*** (0.008)	−0.015** (0.006)	−0.015** (0.006)	0.001 (0.001)	0.000	29
CF_t	−0.004 (0.008)	0.358*** (0.017)	0.038*** (0.014)	0.0001 (0.001)	0.271	1595
ACC_t	0.005 (0.007)	−0.111*** (0.014)	0.167*** (0.013)	−0.006*** (0.001)	0.166	1227
bm_t	−0.337*** (0.078)	−0.869*** (0.091)	−0.705*** (0.098)	0.386*** (0.024)	0.601	2046

(*continued on next page*)

TABLE 2 (Continued)

Variance Decomposition: Cash Flow-Accrual

VARNRF	*VARNCF*	*VARNAC*	*COVRFCF*	*COVRFAC*	*COVCFAC*	*VARRF*
0.000004***	0.011695***	0.007449***	−0.000058***	−0.000008***	−0.007636***	0.004007***
(0.000000)	(0.000224)	(0.000151)	(0.000001)	(0.000002)	(0.000170)	(0.000111)

*, **, *** Indicate significance levels of 10 percent, 5 percent, and 1 percent, respectively.

Panel A lists the parameter estimates for the preliminary earnings VAR model. The model variables include r_t, the three-day cumulative return centered on the preliminary earnings release date (the first element of the state vector z); log of 1 plus income before extraordinary items scaled by initial period book value of equity, roe_t (the second element); and the log book-to-market ratio, bm_t, (the third element). The parameters in the table correspond to the system:

$$z_{i,t} = Az_{i,t-1} + \eta_{i,t} \text{ and } \Omega = E(\eta_{i,t}\eta'_{i,t}).$$

Two numbers are reported for each parameter. The first number is the panel data firm and fiscal quarter fixed effect point estimate of the parameter. The second number in parentheses is a robust jackknife standard error. Panel A also provides the means of the variance decomposition for the preliminary earnings model where the variances are defined as follows:

1. *VARRP* = Total variance of mean-adjusted returns at preliminary earnings release date = *VARNRP* + *VARNEP* − 2**COVRPE*.
2. *VARNRP* = variance of expected-return news at the preliminary earnings release date.
3. *VARNEP* = variance of earnings news at the preliminary earnings release date.
4. *COVRPE* is the covariance of expected-return news with earnings news at the preliminary earnings release date.

The standard errors (in parentheses) of the variances are computed using the Shao and Rao (1993) jackknife method; the clustering variable is year-quarter.

Panel B lists the parameter estimates for the filing date earnings VAR model. The model is the same as for Panel A except that r_t is the cumulative three-day return centered on the 10-Q/K filing date. Panel B also shows the variance decomposition for the SEC filing date earnings model where the variances are defined as follows:

1. *VARRF* = Total variance of mean-adjusted returns at the 10-Q/K filing date = *VARNRF* + *VARNEF* − 2**COVRFE*.
2. *VARNRF* = variance of expected-return news at the 10-Q/K filing date.
3. *VARNEF* = variance of earnings news at the 10-Q/K filing date.
4. *COVRFE* is the covariance of expected-return news with earnings news at 10-Q/K filing date.

Panel C lists the parameter estimates for the cash flow-accrual VAR model. The model variables include r_t the three-day cumulative return centered on the 10-Q/K filing date (the first element of the state vector z); cash flow from operations scaled by beginning of period book value of equity, CF_t, (the second element); accruals scaled by beginning of period book value of equity, ACC_t, (the third element); and the log book-to-market ratio, bm_t, (the fourth element). In addition, Panel C presents the variance decomposition for the cash flow-accrual model where the variances are defined as follows:

1. *VARRF* = Total variance of mean-adjusted returns at the 10-Q/K filing date = *VARNRF* + *VARNCF* + *VARNAC* − 2**COVRFCF* − 2**COVRFAC* + 2**COVCFAC*
   all variables are computed at the 10-Q/K filing date.
2. *VARNRF* = variance of expected-return news.
3. *VARNCF* = variance of cash flow news.
4. *VARNAC* = variance of accrual news.
5. *COVRFCF* = the covariance of expected-return news with cash flows news.
6. *COVRFAC* = the covariance of expected-return news with accruals news.
7. *COVCFAC* = the covariance of cash flows news with accruals news.

The preliminary earnings variable is positively associated with past returns and past preliminary earnings and negatively associated with the past book-to-market ratio. The book-to-market ratio is negatively related to past returns and past preliminary earnings and positively related to the past book-to-market ratio.

Panel A of Table 2 also shows the variance decomposition of unexpected equity returns at the preliminary earnings release date. The standard errors (in parentheses) of the variances are computed using the Shao and Rao (1993) jackknife method based on year-quarters. The variances and the covariance are statistically significant at less than the 1 percent level. The variance contribution of expected return (risk) news (*VARNRP*) is 0.002 percent and the variance contribution of preliminary earnings news (*VARNEP*) is 0.41 percent. The covariance between expected return (risk) and earnings news is −.015 percent indicating that the information contained in expected return (risk) news is negatively correlated with the information contained in earnings news. The total return variance (*VARRP*) computed as in Equation (5) is 0.5 percent.[12] To assess the relative significance of *VARNRP* and *VARNEP*, one compares the volatility (square root of the tabulated variance) of earnings and expected return (risk) news, 0.06 and 0.005, respectively, with the mean log of (1 plus) absolute cumulative returns around the preliminary earnings date of 0.05.[13]

These results indicate that (preliminary) earnings news drives the volatility of contemporaneous unexpected equity returns more than does expected return (risk) news. Nevertheless, the statistically significant variance (and covariance) contribution of expected return (risk) news suggests that it is important to control for risk changes in the analysis. Overall, these finding are consistent with the ubiquitous evidence that the market responds to earnings surprises at preliminary earnings announcements. Furthermore, these findings indicate that the market responds to expected return (risk) news at preliminary earnings announcements.

Panel B of Table 2 shows the parameter estimates of the VAR earnings model around the SEC filing date, which are similar to the estimates around the preliminary earnings announcement. Panel B also presents the variance decomposition on the SEC filing date. The results are similar to the variance decomposition on the preliminary earnings announcement date, although the magnitudes of the variances and the covariance are smaller. More specifically, the variances and covariance are statistically significant at less than the 1 percent level. The variance contribution of expected return news (*VARNRF*) is 0.0004 percent and the variance contribution of earnings news (*VARNEF*) is 0.16 percent.[14] These results also indicate that earnings news at the filing date drives the volatility of contemporaneous unexpected equity returns more than does expected return (risk) news on the SEC filing date. Again, the covariance between earnings news and expected return news is significant and negative (−0.007 percent). Overall, these findings imply that SEC filings convey value-relevant information about shocks to future earnings and about shocks to future firm risk (as mirrored through changes to expected discount rates).

By itself, the earnings number on the SEC filing date cannot convey additional information beyond the preliminary earnings number unless the market interprets the earnings number in conjunction with additional information included in the SEC filing. To explore the possibility that the significant variance contribution of earnings news on the SEC filing date may be attributed to unexpected cash flows or accruals, we follow Callen and Segal (2004) and estimate a VAR model where net income is separated into accruals and net operating cash flows, and investigate whether the variance contributions of these two components of earnings are significantly different from zero. As with the estimation of earnings news on the SEC filing date, the return variable is the three-day cumulative log equity return centered on the SEC filing date.

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[12] Note that the variance of earnings news can exceed or equal the variance of total return news whenever the covariance between expected return news and earnings news is negative.

[13] See Equation (1). By this calculation, expected return (risk) news is slightly more than 8 percent (.005/.06) of earnings news. Note that the returns and ROE in the VAR system are log-transformed.

[14] The square root of earnings news and discount rate news are 0.04 and 0.002, respectively, comparable to the log of (1 plus) absolute cumulative returns around the filing date of 0.03. Hence, discount rate news is 5 percent of earnings news.

Panel C of Table 2 provides parameter estimates of the cash flow-accrual VAR model around the SEC filing date. We find that current returns are negatively and significantly associated with past returns, past cash flows and past accruals. The cash flow equation indicates that cash flows are significantly and positively associated with past cash flows and past accruals. The accruals equation shows that accruals are significantly and negatively associated with past cash flows and past book-to-market ratios and positively associated with past accruals. Finally, the current book-to-market ratio is significantly and negatively (positively) associated with past returns, cash flows, and accruals (past book-to-market ratio).

Panel C of Table 2 shows the variance decomposition of accruals and net operating cash flows on the SEC filing date. All variances and covariances are significant at less than the 1 percent level. The variance contribution of cash flows (1.16 percent) is significantly greater than the variance contribution of accruals news (0.75 percent), suggesting that cash flow news is more value-relevant than accruals news. The expected return (risk) news variance is significantly smaller in magnitude (0.0004 percent), indicating that accrual and cash flow news are the main drivers of equity returns, rather than expected future discount rates.[15] Thus, Panel C strongly implies that the contribution of earnings to the volatility of unexpected equity returns on the SEC filing date is due at least in part to the additional information conveyed by the accruals-cash flow breakdown.

As a sensitivity analysis, we estimate the variance decomposition around non-event (non-filing) periods. Specifically, we estimate the variance contribution of earnings in the three-day window [−3,−1] where day 0 is either the preliminary earnings announcement date or the SEC filing date. Our results show that the variance contribution is significantly greater during the window [−1,+1] than during the window [−3,−1] for both the preliminary earnings announcement and the SEC filing dates. Thus, our methodology picks out the increase in variance contribution due to the additional firm disclosure.

In an additional sensitivity analysis, we examine the effect of clustering on the results. The evidence in Griffin (2003) and our data indicate that a large proportion of firms file financial statements on the due date—45 (90) days after fiscal quarter (year) end—or on the prior day. Clustering potentially affects information flows. On the one extreme, clustering could effectively reduce the multifirm information source to that of a single firm or, on the other extreme, the cross-sectional dependencies could produce a more informative message for all firms due to information transfer from one firm to another.

To examine whether the results are affected by clustering, we conduct two analyses. First, we estimate the mean and median variance contributions separately (1) for firms that file financial statements either on the day prior to the due date or on the due date and (2) for all other firms. Untabulated results show that the variance contributions are all significant and of similar magnitudes for both groups. Second, we compute for each calendar day the proportion of companies that filed the financial statements for the fiscal quarter. We then compute the mean and median variance contributions separately for calendar days with proportion of filers greater than 5 percent and for all other days. Again, the results indicate that clustering has little impact on the variance contributions.

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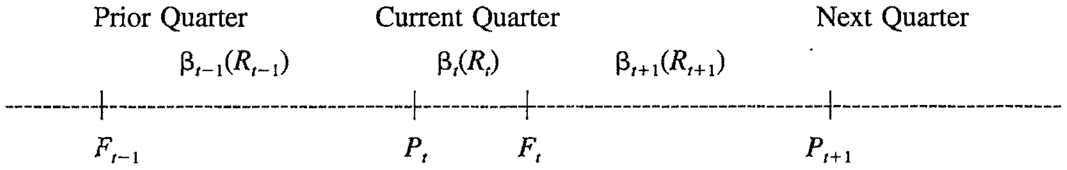
[15] Interestingly, the variance contribution of accruals (*VARNAC*) and cash flows (*VARNCF*) are much larger than the variance contribution of earnings on the SEC filing date (Panel B, Table 2). Part of the difference is certainly attributable to the high correlation between accruals and cash flows news (*COVCFAC* = −0.76%). Nevertheless, one cannot compare directly the variances of accruals and cash flows with the variance of earnings because they are obtained from different models.

Expected Return News, Risk, and Discount Rates

This study maintains that expected return news is a measure of change to firm risk and to firm discount rates. Table 3 tests this maintained assumption formally by regressing changes in beta and changes in *ex post* returns on expected return news. Since expected return news and earnings news are correlated, we also control for earnings news in the regressions. Panel A of Table 3 provides summary statistics of the betas and average daily returns for various time periods. These betas and average daily returns are best described by the schema below. Let F denote the filing date and P the preliminary earnings announcement date. The focus is the change in beta and the average daily return arising out of the preliminary announcement (P_t) and the SEC filing (F_t) in the current quarter t. β_{t-1} (R_{t-1}) denotes the beta (average daily return) from the filing date of the prior quarter to the preliminary date of the current quarter. β_t (R_t) denotes the beta (average daily return) from the preliminary announcement date of the current quarter to the filing date of the current quarter.[16] β_{t+1} (R_{t+1}) denotes the beta (average daily return) from the filing date of the current quarter to the preliminary date of the next quarter. $[\beta_{t+1} + \beta_t]$ ($[R_{t+1} + R_t]$) denotes the beta (average daily return) from the preliminary announcement date of the current quarter to the preliminary announcement date of the next quarter. $[\beta_t + \beta_{t-1}]$ ($[R_t + R_{t-1}]$) denotes the beta (average daily return) from the filing date of the prior quarter to the filing date of the current quarter.



Panel A of Table 3 shows that the betas are relatively stable. The mean (median) of the betas is around 0.69 (0.61). The standard deviation of β_t is much higher than the standard deviation of the betas of the other periods, most probably because there are fewer days between the preliminary announcement and the SEC filing in period t with which to estimate beta as compared to the other periods. In fact, the median number of trading days available to estimate β_t is 13, suggesting that the estimates of β_t are not statistically reliable. The mean (median) average daily returns are from 0.02 to .06 (.01 to .04) depending on the time period. By comparison to the betas, average daily returns are highly volatile, relative to the mean.

Table 3, Panel B shows regressions of *changes* in various betas on expected return (risk) news and earnings news.[17] We first focus on the current quarter preliminary announcement (P_t). The first column of Table 3, Panel B shows that the change in beta from

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[16] All betas are computed after eliminating the three-day period around the event date, be it the preliminary earnings announcement or the SEC filing dates.

[17] Although the correct focus of the analysis is on *changes* in returns, we also find that expected return news is positively correlated with returns after the preliminary announcement and after SEC filing dates. The latter results are consistent with a changes specification if the expected returns in the base periods are cross-sectionally constant.

TABLE 3
Analysis of Changes in Beta, Changes in Returns, and Expected Return News

Panel A: Summary Statistics of the Betas and Averaged Daily Returns

	n	Mean	Std	Q1	Median	Q3
β_{t-1}	52,707	0.697	0.756	0.218	0.608	1.082
β_t	56,016	0.665	1.134	0.060	0.596	1.235
β_{t+1}	51,274	0.696	0.733	0.222	0.610	1.078
$\beta_t + \beta_{t-1}$	55,502	0.699	0.699	0.250	0.615	1.060
$\beta_{t+1} + \beta_t$	54,460	0.697	0.686	0.252	0.611	1.049
R_{t-1}	53,352	0.032	0.392	−0.188	0.022	0.237
R_t	53,367	0.055	0.669	−0.306	0.036	0.405
R_{t+1}	53,339	0.020	0.389	−0.203	0.011	0.229
$R_t + R_{t-1}$	53,333	0.034	0.333	−0.160	0.023	0.209
$R_{t+1} + R_t$	53,344	0.045	0.324	−0.141	0.031	0.215

Panel B: Regression of Changes in Beta on Expected Return News

Dependent Variable	$\beta_t - \beta_{t-1}$	$(\beta_{t+1} + \beta_t) - \beta_{t-1}$	$\beta_{t+1} - \beta_t$	$\beta_{t+1} - \beta_{t-1}$	$\beta_{t+1} - (\beta_t + \beta_{t-1})$
Constant	−0.035*** (0.005)	−0.001 (0.003)	0.035*** (0.005)	0.001 (0.003)	−0.001 (0.003)
Nrp	1.489 (1.292)	2.147*** (0.798)			
Nep	0.424*** (0.095)	0.195*** (0.058)			
Nrf			8.032* (4.186)	7.768*** (2.851)	10.121*** (2.542)
Nef			0.184 (0.211)	0.269* (0.144)	0.344*** (0.128)
n	50773	52675	51201	51244	50773
R^2	0.0004	0.0003	0.0001	0.0001	0.0001

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*, **, *** Indicate significance levels of 10 percent, 5 percent, and 1 percent, respectively.

Panel A provides summary statistics for the market betas and the averaged daily returns. β_{t-1} (R_{t-1}) is the beta (average daily return) from the prior quarter filing date to the current quarter preliminary earnings announcement. β_t (R_t) is the beta (average daily return) from the current quarter preliminary earnings announcement to the current quarter filing date. β_{t+1} (R_{t+1}) is the beta (average daily return) from the current quarter filing date to the next quarter preliminary earnings announcement. $\beta_t + \beta_{t-1}$ ($R_t + R_{t-1}$) is the beta (average daily return) from the prior quarter filing date to the current quarter filing date. $\beta_{t+1} + \beta_t$ ($R_{t+1} + R_t$) is the beta (average daily return) from the current quarter preliminary earnings announcement date to the next quarter preliminary earnings announcement date. The betas are estimated using the market model and daily returns. We use the value-weighted market return as the independent variable.

Panel B shows the regressions results of changes in beta on expected return (risk) news controlling for earnings news. *Nrp* (*Nrf*) is expected return (risk) news on the preliminary earnings (SEC filing) date. *Nep* (*Nef*) is earnings news on the preliminary earnings (SEC filing) date. The dependent variables are the differences in betas across the various periods described above. For example, $\beta_t - \beta_{t-1}$ is the beta from the current quarter preliminary earnings date to the current quarter filing date less the beta from the prior quarter filing date to the current quarter preliminary earnings date.

the period prior to the preliminary announcement to the period after the preliminary announcement but prior to the filing $[\beta_t - \beta_{t-1}]$ is unrelated to unexpected return news but significantly positively related to earnings news. One possible explanation is the noise in

β_t. Alternatively, the second column of Table 3, Panel B shows the regression where the dependent variable is the change in beta from the period prior to the current preliminary announcement to the period after the current preliminary announcement up to the next preliminary announcement $[(\beta_{t+1} + \beta_t) - \beta_{t-1}]$. Here both expected return (risk) and earnings news are highly significant.[18]

The filing date results are more conclusive. The third column of Table 3 shows that the change in beta from the period prior to filing date (but after the preliminary earnings announcement date) to the period after the filing date $[\beta_{t+1} - \beta_t]$ is significantly related to expected return news (at the 10 percent level) but unrelated to earnings news. However, this result too may be a consequence of the high estimation error in β_t. The fourth column of Panel B yields a more meaningful comparison. The dependent variables are the beta after the filing date and before the preliminary announcement in the next quarter less the same beta in the prior quarter $[\beta_{t+1} - \beta_{t-1}]$. The results show that the change in betas is positively and significantly associated (less than the 1 percent level) with expected return (risk) news. The coefficient on earnings news is also positive and marginally significant (at the 10 percent level). The last column of Panel B shows the regression where the dependent variable is the difference between the beta after the filing and before the preliminary announcement in the next quarter and the beta from the prior quarter filing date to the current quarter filing date $[\beta_{t+1} - (\beta_t + \beta_{t-1})]$. The coefficients on both expected return news and earnings news are positive and significant (at less than the 1 percent level). This result is also subject to estimation error since the initial base includes both the preliminary announcement and the filing. Nonetheless, overall these results confirm that expected return news measures changes in risk, especially at the SEC filing date. The above results are robust when *ex post* returns (R_t) instead of betas are used to measure changes in firm discount rates.

Properties of Estimated News Items

This section briefly analyzes (untabulated) the properties of the estimated news items over time, pre- and post-EDGAR, ranked according to earnings quality, and for 10-Ks compared with 10-Qs. To facilitate the data analysis, we generate firm-quarter estimates of earnings news, expected return news, and related variance-covariance contributions by estimating the firm-quarter covariance matrix and assuming that all observations have the same VAR coefficient matrix.[19] This procedure is executed using the VAR coefficient matrix with equity returns as of the preliminary earnings announcement date and repeated using the VAR coefficient matrix with equity returns as of the SEC filing date. A similar approach is used to estimate the variance decomposition of cash flow news and accrual news (Equation (10)), with equity returns as of the SEC filing date.

We find that the variances of expected return news around the preliminary announcement and SEC filing dates are fairly stable over the 13 sample years. The variances of earnings, cash flows, and accrual news are less stable; the variances of earnings news at the preliminary announcement and SEC filing dates increase from 1998 to 2002 and then decrease back to their former level.[20] Cash flow and accrual news volatilities are relatively stable from 1991 to 2001 and then decrease during 2002 and 2003.

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[18] However, this result suffers from the potential contamination of the preliminary announcement date information with the SEC filing date information (in the current quarter).

[19] For example, earnings news (Equation (A4) in the Appendix) is a function of the VAR coefficient matrix (A) and the residuals from the VAR regressions (Equations (12a) through (12c)). Thus, earnings news and the variance contribution of earnings news can be estimated at the firm-quarter level using the VAR coefficient matrix and the vector of residuals $\Sigma_{it} = [e_{1it}, e_{2it}, e_{3it}]$, where e_j is the estimated residual from model j and i (t) is the observation (time) index. The variance-covariance matrix Ω_{it} is computed as $\Sigma_{it}'\Sigma_{it}$.

Comparison of pre- and post-EDGAR variances yields inconclusive results. Whereas the variance contribution of earnings news post-EDGAR is significantly higher than earnings news pre-EDGAR, the variance contribution of cash flow and accrual news is significantly lower post-EDGAR. The lower variance contribution of accruals and cash flows post-EDGAR is attributable to the lower variance contributions of these factors in 2002 and 2003. Excluding these years from the analysis, we find that the variance contributions are of similar magnitude across the two periods.

Investors in firms with a large proportion of accruals are more likely to carefully assess the accrual information provided in SEC filings than investors in firms with a small proportion of accruals for whom the preliminary earnings disclosure may be almost completely sufficient. To examine the relation between the magnitude of accruals and the variances, we average and then rank the proportion of absolute value of accruals to total assets over the prior eight quarters, and assign the observations to quintile portfolios. We find that the average earnings news variance at the preliminary and filing dates as well as the average accruals and cash flow news variances increase monotonically with the proportion of accruals, consistent with our conjecture.[21]

Finally, we find that while 10-Qs yield a significantly higher (at less than the 1 percent level) earnings news volatility relative to 10-Ks for preliminary earnings announcements, the reverse holds for the filing date. Consistent with Griffin (2003), 10-Ks yield a significantly higher (at less than the 1 percent level) earnings news volatility for earnings, accruals, and cash flow news than 10-Qs at the filing date.[22]

Variance Decomposition and the Information Environment

We investigate the association between the variance contribution on the SEC filing date of expected return (risk), accrual and cash flow information components, and our main proxy for information asymmetry, the degree of investor sophistication, by regressing the variance contributions of these components on institutional ownership, after controlling for the information environment of the firm and other variables likely correlated with institutional holdings.[23,24]

We use the following variables as proxies for the firm's information environment or as more general control variables in the regressions below. We include a dummy variable (*EDGAR*) with 1 if the period is post-EDGAR, and 0 otherwise. Given the conflicting findings of the univariate analysis we do not predict the sign of *EDGAR*. A dummy variable

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[20] These results are consistent with excess volatility for 2000 and 2001 shown in Griffin (2003, Figure 1).

[21] Earnings persistence cannot be used to directly classify firms into quintiles to determine the differential variance contribution in each quintile since the VAR model underlying the variance decomposition methodology controls directly for persistence. On this issue, see Callen (2004).

[22] This is likely due the fact that the fourth quarter exhibits increased activity relative to the other quarters on dimensions such as capital expenditures, bookings of accruals, and auditor adjustments. See for example, Callen et al. (1996) on capital expenditures, Das and Shroff (2002) on accruals, and Ettredge et al. (2000) on auditor adjustments. In addition, the 10-K report is audited so that 10-Ks should yield greater earnings news volatility and, hence, unexpected return volatility than 10-Qs.

[23] We repeat all the analyses using the variance contributions of expected return (risk) and earning news from the earnings model at the filing date. The results are similar to those reported.

[24] Following the extant literature, we employ two empirical proxies for the level of investor sophistication: the percentage of shares outstanding held by institutions and the number of institutions holding the shares. Since both proxies yield very similar results, we report our results only for the percentage of shares outstanding held by institutions.

(*FD*), taking on a value of 1 if the period is post-Regulation FD, and 0 otherwise, is included to control for the possibility that earnings news on the SEC filing date changed after Regulation FD when firms are less likely to engage in partial disclosure to select investors.[25] We incorporate a dummy variable, *ANN*, which equals 1 if the report is annual (10-K), and 0 otherwise (if 10-Q). We also include a dummy variable taking the value of 1 if the firm provides balance sheet information on accounts receivable, inventory, and accounts payable in the preliminary earnings release, and 0 otherwise (*PREBS*). Firms that provide more balance sheet information with preliminary earnings should show less cash flow/accrual news volatilities on the SEC filing date.[26]

Similarly, we include a dummy variable taking the value of 1 if the firm provides net operating cash flow information in the preliminary earnings release, and 0 otherwise (*PREOCF*). Holding everything else constant, we expect less accruals/cash flow volatilities on the SEC filing dates if accruals and cash flow information are available in the preliminary earnings release prior to the SEC filing date.[27] The information environment of the firm, specifically its level of information asymmetry, may also be related to company losses because the relation between losses and future cash flows is weaker for loss than for profit firms. In addition, Ertimur (2004) shows that the bid-ask spread for loss firms is greater than for profit firms, indicative of greater information asymmetry by loss firms. To capture the effect of losses on information asymmetry, we incorporate a loss variable (*LOSS*), which is defined as the number of loss quarters in the past 20 quarters.[28] Firms with a greater string of losses are likely to exhibit higher news volatilities on SEC filing dates. We include a variable *ACC_PROP* defined to be the absolute value of the ratio of total accruals to average total assets in the past eight quarters to proxy for the relative importance of accruals in earnings. Finally, we control for the variance contribution of preliminary earnings (*VARNEP*) in the accruals and cash flows regressions since the magnitude of the information content of the 10-Q/-K at the filing date may be correlated with shocks to earnings at the preliminary earning announcement date. We make no prediction about the relationship between *VARNEP* and the volatility of news items at the SEC filing date. Similarly, in the expected return (risk) news regression we include the variance contribution of expected return (risk) news at the preliminary earnings date (*VARNRP*).[29]

Since there is potential endogeneity between institutional investors' investment decisions and the usefulness of financial statement information around SEC filing dates, we estimate the relation between the variance contributions of the earnings components and institutional holdings using two-stage least squares (2SLS) in conjunction with firm and fiscal-quarter fixed effects panel techniques. Table 4 shows the second stage of the 2SLS

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[25] Although Regulation FD limits the firm's ability to disclose information to selected group of investors, it is still an open empirical question whether the total disclosure following Regulation FD has decreased or increased. Consequently, we do not predict the sign of the coefficient on FD.

[26] It should be noted that *PREBS* only measures whether the firm provided information on selected current accruals in the preliminary earnings release, and not about the availability of net operating cash flows and total accruals.

[27] Levi (2004) and Louis et al. (2005) provide evidence that market reactions around preliminary earnings announcements and SEC filing dates are different for firms that disclose sufficient balance sheet information to estimate accruals in their preliminary earnings releases.

[28] See also Huddart and Ke (2004).

[29] Firm size (*SIZE*) is also an information environment proxy because larger firms tend to be followed and monitored by more analysts and investors and, thus, there are likely less information asymmetries between large firms and investors. In addition, there is a stronger market reaction at earnings announcements for small firms (Bamber 1987) suggesting a greater volatility of returns. We do not include *SIZE* in the main regression because of the high correlation with institutional ownership. The results of a regression with *SIZE* instead of institutional ownership are similar to those reported. For a similar reason, we do not include the number of analysts following the firm.

TABLE 4
Regressions of the Variance Contribution of Expected Return (Risk) News, Cash Flow News, and Accrual News on Aggregated Institutional Holdings

	Expected Sign	*VARNRAC*	*VARNCF*	*VARNAC*
Constant	?	0.0076*	1.2225***	1.2618***
		(0.0041)	(0.1565)	(0.0995)
INS_HOL	−	0.0341***	−1.5149***	−2.0545***
		(0.0090)	(0.3384)	(0.2152)
EDGAR	+	0.0020**	0.0911**	0.1029***
		(0.0010)	(0.0359)	(0.0228)
FD	?	0.0004	0.0432	0.1325***
		(0.0012)	(0.0439)	(0.0279)
ANN	+	0.0067***	0.2210***	0.2654***
		(0.0009)	(0.0348)	(0.0222)
ACC_PROP	+	0.0109	9.3509***	2.1836***
		(0.0272)	(1.0251)	(0.6519)
PREBS	−	−0.0001	−0.0668	−0.0228
		(0.0011)	(0.0429)	(0.0273)
PREOCF	−	−0.0047***	0.0084	−0.0139
		(0.0016)	(0.0607)	(0.0386)
LOSS	+	0.0014***	0.0656***	0.0500***
		(0.0002)	(0.0070)	(0.0045)
VARNEP	?		0.0575***	0.0366***
			(0.0165)	(0.0105)
VARNRP	?	0.0588***		
		(0.0006)		

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*, **, *** Indicate significance levels of 10 percent, 5 percent, and 1 percent, respectively.

This table shows the estimated coefficients and related standard errors (in parentheses) for three regressions. The dependent variables are the variance contribution of expected return (risk) news (*VARNRAC*), the variance contribution of accrual news (*VARNAC*), and the variance contribution of cash flow news (*VARNCF*). *INS_HOL* is institutional ownership. All other independent variables are defined in the notes to Table 1 and 2.

The regressions are estimated using two-stage least squares with firm and fiscal quarter fixed-effect panel data to control for the endogeneity in the investment decision of institutional investors. The first stage includes regression of *INS_HOL* on size (log of market value of equity at the beginning of the fiscal quarter), *ROE* (net income divided by beginning book value of equity), Standard & Poor's stock rating, *SP500* (indicator variable that takes the value of 1 if the firm is included in the S&P 500, and 0 otherwise), volume (log of average monthly volume divided by shares outstanding over prior year), dividend yield, leverage (measured as debt to assets), market-adjusted returns over the fiscal year, loss (the proportion of loss quarters to total quarters in the past five years), average sales growth over the prior eight quarters, and R&D (R&D expense divided by sales). We require valid institutional ownership data as well as data on all variables. As a result the sample size is reduced to 55,653 firm quarters.

regression of the variance contributions of expected return (risk), accrual, and cash flow news as of the SEC filing date on the percentage of ownership by institutional investors (*INS_HOL*), and on the other informational and control variables. Following Bushee (2001), in the first stage (nontabulated), institutional ownership is regressed on *SIZE* (log of market value of equity at the beginning of the fiscal quarter), *ROE* (net income divided by beginning book value of equity), Standard & Poor's stock rating, *SP500* (an indicator variable that takes the value of 1 if the firm is included in the S&P 500, and 0 otherwise), volume (log of average monthly volume divided by shares outstanding over prior year), dividend yield, leverage (measured as debt-to-assets), market-adjusted returns over the fiscal year, loss (the proportion of loss quarters to total quarters in the past five years), average sales growth over the prior 8 quarters, and R&D (R&D expense divided by sales).

Consistent with H4, Table 4 shows that the variance contributions of accrual news and cash flow news are negatively and significantly (at less than the 1 percent level) related to the level of investor sophistication as proxied by institutional holdings. Contrary to our expectation the coefficient on *INS_HOL* is positive and significant in the expected return (risk) regression. However, as we show below, this result is attributable to the aggregation of institutional ownership across investor types. The coefficient on *EDGAR* is positive and significant suggesting that volatilities have increased post-EDGAR. The coefficient on *ANN* is positive and significant, indicating larger return volatility around annual reports than quarterly reports.[30] As expected, firms with a higher proportion of accruals display larger accrual and cash flows volatilities. Similarly, firms with more extensive losses also show significantly larger news volatilities, consistent with loss firms having greater information asymmetry. In addition, firms with higher earnings news volatilities on the preliminary earnings date also display higher news volatilities on the filing date, indicating that the informativeness of the financial report is positively related to the initial shock to earnings. The *FD* variable is positive and statistically significant in the cash flows regression only. Although *PREBS* and *PREOCF* have the expected sign, they are not statistically significant (*PREOCF* is negative and significant in the expected return variance regression).

Table 5 replicates Table 4 except that institutional holdings are disaggregated into "dedicated," "quasi-indexer," and "transient" investors as per Bushee (1998, 2001). The results are similar to those reported in the baseline regressions and Table 4, so only the main differences are discussed. The coefficient on *FD* is now negative and significant across the three regressions, and the coefficients on *EDGAR* are now insignificant (with the exception of the cash flows regression). Most importantly, consistent with H5, while there is a significantly negative relationship between the variance contributions of expected return (risk), cash flows, and accruals news components and institutional holdings by "dedicated" investors and "quasi-indexers," there is a significantly positive relationship between the variance contributions of the expected return (risk) and cash flow news components and institutional holdings by "transient" investors. These results are consistent with longer-term, stable investors having the incentive to acquire more information about the firm prior to the filing date in order to better monitor firm performance and their investments. In addition, the non-negative coefficient on transient holdings is consistent with "transient" (momentum) investors focusing on short-term gains by investing in smaller more risky firms and minimizing their investment in information gathering and analysis.

Overall, the results in Table 4 and Table 5 indicate that the variance contributions of the three informational components are positively associated with the level of information asymmetry between the firm and its investors. The greater the information asymmetry the greater is the news contained in the financial statements and, therefore, the greater is the impact of the financial report on the volatility of unexpected stock returns.

VI. CONCLUSION

This study analyzes informational components that are value-relevant on (the preliminary announcement and) the SEC filing date using a variance decomposition analysis. Value relevance of an informational component is measured by its contribution to the volatility of unexpected returns. Although the variance decomposition methodology has many

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[30] Note that *ANN* represents the fiscal fourth quarter. The coefficients of the second and third fiscal quarter are suppressed in the tables.

TABLE 5
Regressions of the Variance Contribution of Expected Return (Risk) News, Cash Flow News, and Accrual News on Disaggregated Institutional Holdings

	Expected Sign	*VARNRAC*	*VARNCF*	*VARNAC*
Constant		0.0631***	7.7924***	3.8247***
		(0.0133)	(0.6849)	(0.3643)
DED_HOL	−	−0.1319	−38.5257***	−15.4478***
		(0.0876)	(4.5196)	(2.4044)
TRA_HOL	+	0.1215***	4.0243***	0.4436
		(0.0148)	(0.7558)	(0.4021)
QIX_HOL	−	−0.1567***	−12.0421***	−6.9355***
		(0.0269)	(1.3636)	(0.7254)
EDGAR	+	0.0000	−0.1284**	0.0105
		(0.0010)	(0.0533)	(0.0284)
FD	?	−0.0077**	−1.2993***	−0.3725***
		(0.0031)	(0.1612)	(0.0857)
ANN	+	0.0057***	0.2305***	0.2570***
		(0.0010)	(0.0502)	(0.0267)
ACC_PROP	+	−0.0191	5.3151***	0.6339
		(0.0288)	(1.4714)	(0.7828)
PREBS	−	0.0004	−0.1329**	−0.0410
		(0.0012)	(0.0612)	(0.0326)
PREOCF	−	−0.0026	0.0755	0.0185
		(0.0017)	(0.0849)	(0.0452)
LOSS	+	0.0011***	0.0368***	0.0377***
		(0.0002)	(0.0101)	(0.0054)
VARNEP	?		−0.0182	0.0095
			(0.0246)	(0.0131)
VARNRP	?	0.0581***		
		(0.0007)		

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*, **, *** Indicate significance levels of 10 percent, 5 percent, and 1 percent, respectively.

This table replicates Table 4 where institutional ownership (*INS_HOL*) is disaggregated by the proportion of shares held by "dedicated" institutional investors (*DED_HOL*), "quasi-indexer" institutional investors (*QIX_HOL*), and "transient" institutional investors (*TRA_HOL*). All other variables are as defined in Table 4.

The regressions are estimated using two-stage least squares for firm and fiscal quarter fixed-effect panel data to control for the endogeneity in the investment decision of institutional investors (see the notes to Table 4).

strengths, it also has limitations. These include the fact that variance contributions cannot be compared across models, so it is impossible to assess, for example, whether the decomposition of earnings into cash flows and accruals actually provides more explanatory power than earnings alone for the volatility of unexpected returns. An additional limitation is that when too many variables are included in the VAR system, the estimated coefficients tend to become unstable. Thus, the researcher has to trade off a comprehensive set of possible determinants of the volatility of unexpected returns for parsimony. As a result, the model is best suited to assess the relative volatility contributions of different pre-determined earnings and expected return components and less useful in determining which particular components (among many alternatives) might explain volatility.

Subject to these limitations, this study finds that both expected return (risk) and earnings news are value-relevant on the preliminary earnings and SEC filing dates with earnings news the dominant informational component in terms of economic significance. Expected return (risk) news, accrual news, and cash flow news are all found to be value-relevant with both accrual news and cash flows dominating expected return news in terms of economic significance. Contrary to Callen and Segal (2004) for annual data, cash flow news is found to be more value-relevant than accrual news in driving the volatility of unexpected returns around the SEC filing date. Furthermore, this study shows that the earnings, cash flow, and accruals news exhibit changes in time-series pattern, especially in later years. Consistent with recent research, this study also shows that earnings, cash flow, and accruals news are more value-relevant at the SEC filing date for 10-K forms than for 10-Q forms. They are also more value-relevant for firms with a larger proportion of accruals in earnings. The results comparing pre- and post-EDGAR are inconclusive. Whereas the variance contribution of earnings news post-EDGAR is significantly higher than earnings news pre-EDGAR, the variance contribution of cash flows and accruals news are significantly lower post-EDGAR.

Although earnings news is more economically significant than expected return (risk) news in driving unexpected return volatility at both the preliminary and SEC filing dates, expected return (risk) news is, nevertheless, shown to be significantly correlated with *changes* in betas and returns at the preliminary and SEC filing dates, indicating that expected return news is associated with changes in firm risk and discount rates.

This study also documents that each of the expected return (risk), accrual, and earnings news contain less value-relevant information at the SEC filing date for firms that have a higher proportion of sophisticated investors with long-term investment horizons. The opposite result obtains for transient momentum investors. This is consistent with the notion that sophisticated investors with long-term investment horizons are likely to gather and process information that helps them to assess the quality of preliminary earnings before the SEC filings become available. In contrast, transient momentum investors are unlikely to gather information prior to the SEC filing and, if anything, they are more likely to invest in firms with excess volatility. These findings support a negative association between the value relevance of SEC filings and the degree of information asymmetry between the firm and its investors.

APPENDIX

Variance Decomposition Estimation

As shown by Campbell (1991), the variance decomposition can be implemented empirically by combining the residuals from the VAR estimation with the unexpected current return valuation equation (Equation (1)). Formally, let $ei' = (0,\ldots,1,\ldots,0)$, where the 1 is in the *i*th position. The unexpected change in returns is computed as:

$$r_t - E_{t-1}(r_t) = e1'\eta_{1t}. \tag{A1}$$

Equation (11) implies that forecasts of the state vector $z_{i,t}$ can be computed as:

$$E_t[z_{i,t+1+j}] = A^{j+1}z_{i,t}. \tag{A2}$$

Using Equation (A1), the revision in expected future returns (expected return or risk news) is computed as:

$$\Delta E_t \sum_{j=1}^{\infty} \rho^j r_{t+j} = E_t \sum_{j=1}^{\infty} \rho^j r_{t+j} - E_{t-1} \sum_{j=1}^{\infty} \rho^j r_{t+j} = e1'\rho(I - \rho A)^{-1}\eta_{i,t} = \lambda_1 \eta_{i,t}. \tag{A3}$$

Similarly, the revision in expected future earnings, earnings news, is computed residually as:[31]

$$\Delta E_t \sum_{j=0}^{\infty} \rho^j roe_{t+j} \equiv r_t - E_{t-1} r_t + \Delta E_t \sum_{j=1}^{\infty} \rho^j r_{t+j} = (e1 + \lambda_1)\eta_{i,t}. \tag{A4}$$

The variances and covariance of the variance decomposition (Equation (5)) are obtained from the estimated variance-covariance matrix $\Omega = E(\eta_{i,t}, \eta'_{i,t})$. Specifically, the variances and covariance of expected return news and earnings news are computed as:

$$\text{var}(Nr_t) = \lambda'_1\, \Omega\, \lambda_1;\ \text{var}(Ne_t) = \lambda'_2\, \Omega\, \lambda_2;\ \text{cov}(Nr_t, Ne_t) = \lambda'_1\, \Omega\, \lambda_2.$$

REFERENCES

Badrinath, S. G., and S. Wahal. 2002. Momentum trading by institutions. *Journal of Finance* 57 (6): 2449–2478.

Balsam, S., E. Bartov, and C. Marquardt. 2002. Accruals management, investor sophistication, and equity valuation: Evidence form 10-Q filings. *Journal of Accounting Research* 40 (4): 987–1012.

Bamber, L. S. 1987. Unexpected earnings, firm size, and trading volume around quarterly earnings announcements. *The Accounting Review* 62 (3): 510–532.

Barth, M. E., W. E. Beaver, J. R. M. Hand, and W. R. Landsman. 1999. Accruals, cash flows, and equity values. *Review of Accounting Studies* 4 (3/4): 205–229.

Bennett, J., R. Sias, and L. Starks. 2000. Investment competition, greener pastures and dynamic institutional preferences. Working paper, The University of Texas at Austin.

Bernard, V. L., and T. L. Stober. 1989. The nature and amount of information in cash flows and accruals. *The Accounting Review* 64 (4): 624–652.

Bowen, R. M., D. Burgstahler, and L. A. Daley. 1987. The incremental information content of accrual versus cash flows. *The Accounting Review* 62 (4): 723–747.

Bushee, B. 1998. The influence of institutional investors on myopic R&D investment behavior. *The Accounting Review* 73 (3): 305–333.

———. 2001. Do institutional investors prefer near-term earnings over long-run value? *Contemporary Accounting Research* 18 (2): 207–246.

Callen, J. L., J. Livnat, and S. G. Ryan. 1996. Capital expenditures: Value-relevance and fourth-quarter effects. *Journal of Financial Statement Analysis* 1 (3): 13–24.

———. 2004. Shocks to shocks: A theoretical foundation for the information content of earnings. Working paper, University of Toronto.

———, and D. Segal. 2004. Do accruals drive firm-level stock returns? *Journal of Accounting Research* 42 (3): 527–560.

———, O. K. Hope, and D. Segal. 2005. Domestic and foreign earnings, stock return variability, and the impact of investor sophistication. *Journal of Accounting Research* 43 (3): 377–412.

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[31] Following Vuolteenaho (2002), Callen and Segal (2004), and Callen et al. (2005), we assume that $\rho = 0.967$ on annual basis. Since we use quarterly date we assume that $\rho = 0.99$. Following the literature, earnings news is computed on a residual basis as in Equation (A4). In order that accrual and cash flow news are measured comparably, we compute both on a nonresidual basis.

Campbell, J. Y. 1991. A variance decomposition for stock returns. *Economic Journal* 101 (405): 157–179.

———, and J. Ammer. 1993. What moves the stock and bond markets? A variance decomposition for long-term asset returns. *Journal of Finance* 48 (1): 3–37.

———, A. W. Lo, and A. C. MacKinlay. 1997. *The Econometrics of Financial Markets.* Princeton, NJ: Princeton University Press.

Chung, K., R. Jacob, and Y. B. Tang. 2003. Earnings management by firms announcing earnings after SEC filing. *International Advances in Economic Research* 9 (2): 152–162.

Cochrane, J. H. 2001. *Asset Pricing.* Princeton, NJ: Princeton University Press.

Cready, W., and P. Mynatt. 1991. The information content of annual reports: A price and trading response analysis. *The Accounting Review* 66 (2): 291–312.

Das, S., and P. K. Shroff. 2002. Fourth quarter reversals in earnings changes and earnings management. Working paper, University of Illinois at Chicago and University of Minnesota.

Dechow, P. M. 1994. Accounting earnings and cash flows as measures of firm performance: The role of accounting accruals. *Journal of Accounting and Economics* 18 (1): 3–42.

Easton, P. D., and M. E. Zmijewski. 1993. SEC form 10K/10Q reports and annual reports to shareholders: Reporting lags and squared market model prediction errors. *Journal of Accounting Research* 31 (1): 113–129.

Ertimur, Y. 2004. Accounting numbers and information asymmetry: Evidence from loss firms. Working paper, Stanford University.

Ettredge, M. L., D. T. Simon, D. B. Smith, and M. S. Stone. 2000. The effect of the external accountant's review on the timing of adjustments to quarterly earnings. *Journal of Accounting Research* 38: (1): 195–207.

Foster, T., and D. Vickrey. 1978. The incremental information content of the 10-K. *The Accounting Review* 53 (4): 921–934.

———, D. Jenkins, and D. Vickrey. 1983. Additional evidence on the incremental information content of the 10-K. *Journal of Business Finance and Accounting* 10 (1): 57–66.

———, ———, and ———. 1986. The incremental information content of the annual report. *Accounting Business and Research* 16 (62): 91–98.

Griffin, P. 2003. Got information? Investor response to form 10-K and form 10-Q EDGAR filings. *Review of Accounting Studies* 8 (4): 433–460.

Healy, P. M., and K. G. Palepu. 2001. Information asymmetry, corporate disclosure, and the capital markets: A review of the empirical disclosure literature. *Journal of Accounting and Economics* 31 (1–3): 405–440.

Huddart, S., and B. Ke. 2004. Information asymmetry and cross-sectional variation in insider trading. Working paper, The Pennsylvania State University.

Kothari, S. P. 2001. Capital markets research in accounting. *Journal of Accounting and Economics* 31 (1–3): 105–231.

Lev, B. 1989. On the usefulness of earnings and earnings research: Lessons and directions from two decades of empirical research. *Journal of Accounting Research* 27 (Supplement): 153–201.

Levi, S. 2004. Voluntary disclosure of accruals in preliminary announcements and the pricing of accruals. Ph.D. dissertation, New York University.

Livnat, J., and P. Zarowin. 1990. The incremental information content of cash-flow components. *Journal of Accounting and Economics* 13 (1): 25–46.

Louis, H., D. Robinson, and A. Sbaraglia. 2005. An integrated explanation for the potential mispricing of discretionary earnings. Working paper, The Pennsylvania State University.

Monks, R., and N. Minow. 1995. *Corporate Governance.* Cambridge, MA: Blackwell Publishing.

Pfeiffer, R. J. Jr., and P. T. Elgers. 1999. Controlling for lagged stock price responses in pricing regressions: An application to the pricing of cash flows and accruals. *Journal of Accounting Research* 37 (1): 239–248.

Porter, M. 1992. *Capital Choices: Changing the Way America Invests in Industry.* Boston, MA: Council on Competitiveness/Harvard Business School.

Qi, D., W. Wu, and I. Haw. 2000. The incremental information content of SEC 10-K reports filed under the EDGAR system. *Journal of Accounting, Auditing, and Finance* 15 (1): 24–46.

Rayburn, J. 1986. The association of operating cash flow and accruals with security returns. *Journal of Accounting Research* 24 (Supplement): 112–133.

Shao, J., and J. N. K. Rao. 1993. Jackknife inference for heteroscedastic linear regression models. *Canadian Journal of Statistics* 21: 377–385.

Sloan, R. G. 1996. Do stock prices fully reflect information in accruals and cash flows about future earnings? *The Accounting Review* 71 (3): 289–315.

———. 1999. Accruals, cash flows, and equity values: Comment. *Review of Accounting Studies* 4 (3/4): 231–234.

Stice, E. 1991. The market reaction to 10-K and 10-Q filings and to subsequent *The Wall Street Journal* earnings announcements. *The Accounting Review* 66 (1): 42–55.

Subramanyam, K. R. 1996. The pricing of discretionary accruals. *Journal of Accounting and Economics* 22 (1–3): 249–281.

Vuolteenaho, T. 2002. What drives firm-level stock returns? *Journal of Finance* 57 (1): 233–264.

Wilson, G. P. 1986. The relative information content of accruals and cash flows: Combined evidence at the earnings announcement and annual report release date. *Journal of Accounting Research* 24 (Supplement): 165–200.

———. 1987. The incremental information content of the accrual and funds components of earnings after controlling for earnings. *The Accounting Review* 62 (2): 293–322.

Xie, H. 2001. The mispricing of abnormal accruals. *The Accounting Review* 76 (3): 357–373.

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Uncertainty, Legal Liability, and Incentive Contracts

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ABSTRACT: To address agents' moral hazard over effort, incentive contracts impose risk on the agents. As performance measures become noisier, the conventional agency analysis predicts that principals will reduce the incentive weights assigned to such measures. However, prior empirical results (Prendergast 2002) frequently find the opposite, i.e., incentive weights are larger (agents bear more risk) in more uncertain environments. This paper provides new evidence on the association between the extent of uncertainty and the level of risk imposed on agents. In the context of contracts between managed care organizations and physicians, we examine the effect of task characteristics and the legal liability environment on the extent of risk that physicians bear. We derive the optimal weighting of multiple performance measures in a model of a physician's choice of revenue-generating and cost-control efforts. The model predicts that physicians who face less task uncertainty bear *more cost risk* in their contracts, as predicted by the conventional moral hazard model. Likewise, the model predicts that as the association between task uncertainty and legal liability uncertainty becomes stronger, physicians bear *less cost risk* in their contracts. Our empirical results generally support these predictions. We offer an explanation for why these results tend to be consistent with the conventional moral hazard analysis, contrary to empirical results in a number of previous studies.

Keywords: *risk-sharing; incentive contracts; legal liability; task characteristics.*

Data Availability: *Data are available from sources identified in the paper.*

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I. INTRODUCTION

Risk-sharing plays an important role in the design of control systems. Firms design incentive contracts that influence agents' private action choices because the agents share the risk of uncertain performance measures that depend on these choices. This paper focuses on two factors associated with the riskiness of performance measures—agents' task characteristics and their legal liability. We examine the influence of cross-sectional variations in these two sources of uncertainty on the extent of risk that agents, who are subject to moral hazard, bear in incentive contracts.

The particular context we analyze is incentive contracting between managed care organizations (MCOs) and physicians. Physicians face uncertainty about a patient's actual medical condition, and therefore uncertainty about the revenue and cost implications of treating the patient, as well as uncertain malpractice liability under alternative treatment protocols. Physicians frequently work under contracts that hold them responsible for some or all of the uncertain cost of care. Likewise, malpractice liability exposes physicians to the risk associated with the uncertain outcomes of patient care combined with an uncertain legal liability process.

To explain existing contractual arrangements, the economics and accounting literatures have adopted an agency theory perspective that emphasizes the trade-offs between the incentive and risk-bearing effects of contractual provisions (Holmstrom 1979). The agency perspective on moral hazard concludes that imposing financial risk on agents can influence them to work harder and consume fewer resources, but simultaneously requires making greater expected payments to the agents. The increased payments compensate risk-averse agents for bearing a portion of the exogenous risk associated with uncertain outcomes. In turn, fixing the pay-performance sensitivity, the required risk premium increases as the extent of exogenous risk increases. Therefore, the traditional agency approach to the moral hazard problem suggests that because risk-sharing incentives become more expensive in more uncertain environments, the principal will reduce the agent's pay-performance sensitivity as the environment becomes riskier.

The preceding agency approach to the design of incentive contracts treats the legal system as a uniform mechanism to enforce contracts between principals and agents. Our analysis extends this perspective by recognizing that the design of incentive contracts potentially depends on the nature and extent of the agents' malpractice liability, which can vary significantly across jurisdictions. An agent's optimal action choice will depend on both the structure of the agent's contract with the principal and also on how alternative action choices affect the agent's legal liability. In turn, the principal will design the agent's contract recognizing how potential legal liability can influence the agent's action choice. In addition, to the extent that the uncertain performance measure and malpractice liability outcomes are correlated, the increase in legal liability may interact with the uncertainty in performance measures, producing a further influence on the optimal contract design.

This paper examines incentive contracts for physicians as a new test of the traditional agency moral hazard analysis. First, we derive hypotheses based on a model of moral hazard with multiple actions and multiple performance measures. The hypotheses relate physicians' task characteristics, which are proxies for the extent of cost uncertainty facing the physician, and uncertain malpractice liability,[1] stemming from state malpractice provisions, to the form of contracts between managed care organizations (MCOs)[2] and physicians.

Next, we test the hypotheses derived from the analytical model, using data from a national survey of physicians. With respect to task characteristics, physician specialists face greater exogenous risk than primary care physicians (PCPs). Hence, the analytical model and conventional moral hazard analysis predict that PCPs will bear more cost risk than specialists. Our empirical results support this prediction. We find that specialists, who face more exogenous task-related risk, bear less cost risk than PCPs. Likewise, our results indicate that physician task characteristics also interact with the nature of physicians' malpractice liability in determining the form of MCO-physician contracts. In particular, in environments in which cost outcomes are strongly associated with legal liability outcomes, physicians experience a further reduction in the cost risk in their contracts with MCOs.

Our results have several implications. First, we add to the growing literature that provides empirical tests of alternative models of optimal contract design. Our results are generally consistent with the traditional moral-hazard-based agency trade-off between incentives and risk-bearing and, therefore, they contrast with empirical results in certain other contexts as described by Prendergast (2002). We discuss how the trade-off between risk-sharing and incentives can change when agents have private information (e.g., Prendergast 2002). Based on this distinction, we explain why the moral hazard model has greater predictive power in our setting than in the environments considered by Prendergast.

Second, we demonstrate the potentially important role that legal institutions can play in the design of incentive contracts. Finally, our results identify an interesting interaction between exogenous uncertainty and legal liability in which a stronger association between legal liability and unfavorable cost control outcomes leads to weaker incentives for physician specialists.

The rest of this paper is organized as follow. Section II discusses previous research and Section III presents the model and hypotheses. Section IV describes the data and Section V presents the results. Section VI provides a discussion and conclusions.

II. BACKGROUND AND RELATED LITERATURE

In addition to agency moral hazard and private information models, this paper relates to three other streams of research in accounting and economics. First, analytical research in economics has addressed the optimal design of MCO-physician contracts when agents face legal liability. Gal-Or (1999) characterizes optimal contracts between MCOs and risk-neutral physicians in a model in which physicians and MCOs share both cost risk and legal liability risk. Consistent with Gal-Or (1999), we assume that physicians incur disutility from their unobservable cost control effort, leading to a moral hazard problem with respect to this effort. Our approach differs from Gal-Or's (1999) in that we assume that physicians are risk-averse and also that every physician faces significant legal liability costs that cannot be shared with the MCO.[3] Our empirical results are consistent with risk-aversion playing an important role in the design of optimal incentive arrangements.

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[1] Although most agents face potential legal liability for their actions, we emphasize legal liability for physicians because we expect them to be particularly sensitive to malpractice liability concerns.

[2] Managed care organizations (MCOs) include Health Maintenance Organizations (HMOs), preferred provider organizations (PPOs), point-of-service (POS) plans, and independent physician associations (IPAs) (Brown and Hartung 1998).

[3] Several factors limit an MCO's ability to share all legal liability costs with physicians. First, a significant portion of such cost is nonfinancial (personal reputation, self-esteem, personal cost of defending against suits, etc.), and hence cannot be shared. Second, an MCO may incur political costs from regulators who disapprove of actions that would shield physicians from the financial costs of negligent care.

Second, prior empirical research in health care settings has documented various features of physicians' incentive arrangements with MCOs. Gold et al. (1995) survey the literature on the form of incentive contracts between MCOs and physicians. They cite results from Hillman (1987) and Hillman et al. (1992), documenting the extent to which MCO-physician contracts vary with the organizational structure of the MCO and the nature of physician specialties. In contrast to our paper's emphasis on testing hypotheses concerning the determinants of the extent of risk that agents bear in incentive contracts, Gold et al. (1995) review and synthesize studies that focus on documenting the various forms of incentive arrangements between MCOs and physicians. They emphasize that in some cases, rather than contracting directly with physicians, MCOs instead contract with physician practices or other intermediaries who, in turn, contract with individual physicians. This produces a three-tier contractual relation (MCO, physician practice, individual physician) rather than the typical two-tier relation that we assume. In such three-tier cases, our study focuses on the contract between the MCO and the physician practice, whereas Ittner et al. (2006) address the contracts between the practices and individual physicians. They find that the two contracts tend to be congruent in the sense that when the MCO-physician practice contract provides stronger cost control incentives for the practice (as measured by greater reliance on capitation), the practice responds by relying less on fee-for-service contracts to strengthen individual physicians' cost control incentives.[4]

Leone (2002) hypothesizes that certain HMO-level factors will be associated with more high-powered cost-control incentives (more capitation) in HMO-PCP contracts. He concludes that HMOs design their contracts with PCPs to minimize agency costs as a function of the HMO's organizational structure (HMO type), customer mix (proportion of Medicare patients), and environment (national versus local HMO). Bazzoli et al. (1999) find that physicians bear more risk (more capitation) when there are more HMOs in the local market, when the HMOs control a greater market share, and in the presence of fewer physician group practices. Our approach is distinguished from the preceding two studies by our use of physician-level rather than HMO-level data. However, based on their findings, we control for the size of physician practices as well as for the relative bargaining power of MCOs and physician practices, as measured by the extent of each practice's reliance on managed care revenue.

Third, empirical research provides evidence concerning the effect of legal liability on physicians' practice patterns. Kessler and McClellan (1996, 2000) document that physicians faced with greater legal liability engage in more defensive medicine[5] by consuming additional resources on tests and procedures that have no real medical benefits for the patient.[6] They also document that legal reforms that reduced physicians' legal liability in some states over the period 1984–1994 significantly reduced healthcare costs. Our analysis differs from Kessler and McClellan (1996, 2000) with respect to the MCO's choice of incentive contracts for physicians. Kessler and McClellan (1996, 2000) treat the MCO's choice of contract design features as exogenous in their analysis. In contrast, we assume that each MCO designs contracts with physicians based on the nature and extent of physicians' legal liability in each state in which the MCO operates, and how that liability is associated with the physician's task characteristics, i.e., PCP versus specialist.

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[4] As explained further in the next section, MCOs contract with physician practices using a combination of capitation and fee-for-service contracts. Capitation contracts impose more risk on physicians because capitation payments to the physician are not adjusted for the nature and extent of actual services provided, whereas payments are adjusted under fee-for-service contracts.

[5] Defensive medicine is a form of moral hazard whereby the physician provides "precautionary medical treatments that have minimal expected medical benefit out of fear of legal liability" (Kessler and McClellan 2000, 1).

[6] In a professional publication for physicians, Davenport (2000) advises physicians on limiting their malpractice liability in the high-risk areas of myocardial infarction, breast cancer, appendicitis, lung cancer, and colon cancer. Davenport (2000) stresses the importance of performing appropriate tests, following up the test results, and most importantly, documenting the nature and basis of care.

III. MODEL AND HYPOTHESES

This section develops our model of incentive contracts between MCOs and physicians.[7] Consistent with the standard moral hazard formulation, the model focuses on the MCO's trade-off between the incentive benefits of imposing risk on physicians and the cost to compensate physicians for bearing the risk. The model demonstrates how this trade-off varies as the physician's environment changes. In particular, the model reflects two determinants of the risk that incentive contracts impose on physicians. The first is the extent of noise in the MCO's measure of the physician's cost control performance. The second is the association between the noise in the cost measure and the noise in the outcome of the malpractice liability system.[8] The model demonstrates how an MCO adjusts the extent of risk borne by physicians as these parameters change.

We analyze the changes in these features of the physician's environment because we believe they reflect key forces that determine the MCO's trade-off of incentive benefits versus the cost of risk-sharing. First, moving from an environment in which primary care physicians (PCPs) provide more standardized, routine care to an environment facing specialists, who provide less standardized care, is likely to produce a significant increase in the extent of exogenous noise in the cost measure. From the model, we derive the prediction that the MCO will impose less cost risk on physicians as the cost measure becomes noisier, i.e., as we move from the typical environment facing PCPs to that facing specialists. Second, for specialists we expect a strong association between unfavorable cost control realizations and large malpractice liability outcomes. The model establishes conditions under which this association leads MCOs to impose less cost risk on specialists when malpractice liability is greater.

We employ a Linear contract, Exponential utility, Normal distribution (LEN) formulation with multiple actions and multiple performance measures, similar to that in Datar et al. (2001). The physician takes two actions (a_1, a_2) and the MCO employs two performance measures (y_1, y_2) in the physician's incentive contract. Action a_1 involves providing patient care and action a_2 is effort to control the cost of that care. The first performance measure (y_1) reflects the physician's productivity in generating revenue, while the second measure (y_2) captures cost containment effort.

Let the MCO's payment to the physician be $\alpha + \beta_1 y_1 + \beta_2 y_2$, where α is fixed salary and β_1 and β_2 are incentive weights. Here, y_1 measures the extent of services performed by the physician and y_2 measures cost control performance. For instance, y_1 could be total billable charges or total procedures performed by the physician and y_2 could be the cost variance, (budgeted cost − actual cost), where the budgeted cost represents a capitation payment (a fixed payment per member per month [PMPM]) from the MCO to the physician.[9] For a physician receiving no payments from fee-for-service (FFS) contracts and all payments from capitation contracts (assuming for illustration that the incentive weights must sum to 1), we would have $\beta_1 = 0$ and $\beta_2 = 1$. Conversely, for a physician receiving all payments from fee-for-service contracts and none from capitation, the incentive weights would be $\beta_1 = 1$ and $\beta_2 = 0$. Most physicians generate compensation from multiple contracts, reflecting a combination of capitation and fee-for-service arrangements. For such physicians, an increase in the ratio β_2/β_1 reflects an increasing reliance on capitation and more cost risk imposed on the physician.

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[7] Our model assumes that MCOs contract directly with individual physicians, but our empirical tests are based on contracts between MCOs and physician practices. For smaller physician practices, which constitute the bulk of our observations (87 percent of our observations are from practices with ten or fewer physicians), incentives for the practice and for individual physicians should be relatively similar. However, incentives may differ more significantly in larger physician practices, which may employ a second set of contracts between the practice and the individual physician. In the absence of data on the contracts between the physician practice and individual physicians, we rely on the assumption that the overall incentives will be similar in both sets of contracts. To the extent that the incentives differ, this should operate against our finding support for the hypotheses derived from our analytical model.

[8] Physicians also face risk with respect to the level of revenue their services will generate. However, because this third source appears to be less important empirically, we focus on the first two sources of risk.

The physician's revenue-generation and cost control efforts also determine the MCO's economic value. The physician's efforts produce value (x) for the MCO as $x = f_1a_1 + f_2a_2 + e_x$, where x is assumed not to be contractible. Further, $f_i > 0$, $i = 1,2$, is the sensitivity of value creation for the MCO to effort i by the physician, and e_x is the impact of all other factors on MCO value creation, where $e_x \sim N(0,\text{var}_x)$. This formulation indicates that physicians increase MCO value by delivering high-quality services (a_1), which attract and retain profitable patients for the MCO, as well as by practicing cost-conscious use of resources (a_2). Performance measures y_1 and y_2 depend on actions a_1 and a_2 as $y_i = g_{i1}a_1 + g_{i2}a_2 + e_i$, $i = 1,2$, where g_{ij} is the sensitivity of performance measure i to action j, and e_i reflects the effect of all other factors on performance measure y_i, where $e_i \sim N(0, \text{var}_i)$, $i = 1,2$. We assume that $g_{11} > 0$, $g_{22} > 0$, and $g_{12} = g_{21} = 0$, which indicate that revenue-generating effort (a_1) increases the revenue measure (y_1) but has no effect on the cost control measure (y_2). Likewise, cost control effort (a_2) improves the cost control measure (y_2) but has no effect on the revenue measure. We also assume that $\text{cov}(e_1, e_2) = 0$, reflecting no association between the revenue and cost measure error terms.

Let the physician have a negative exponential utility function, $U(W) = -e^{-rW}$, where r is the coefficient of absolute risk aversion. The physician's net wealth is $W = w(y_1,y_2) - C(a_1,a_2) - L(a_1,a_2,k)$, where w is the physician's compensation from the MCO, C is the cost to the physician of patient care, and L is the physician's loss from legal malpractice liability.

Let $C(a_1,a_2) = .5a_1^2 + .5a_2^2$ be the monetary equivalent of the physician's disutility from effort. Let $L(a_1,a_2,k) + e_L$ be the loss to the physician from malpractice liability claims, with $0 \leq L < \bar{L}$, where $\bar{L}$ is a finite upper bound on a physician's malpractice liability. We assume that $L\,(.) + C\,(.)$ is convex, and together with the linear objective function, this assumption ensures that the physician's choice of actions is interior.[10] Letting L_i denote the partial derivative of L with respect to argument i, where $i = 1,2,3$, we assume that $L_1 < 0$, indicating that physicians can reduce their expected malpractice loss by providing more extensive care (a_1), such as by performing additional tests and other procedures,[11] and $L_2 > 0$, indicating that more cost control effort by the physician (a_2) increases the physician's expected malpractice exposure. Let $k > 0$ be a parameter reflecting the extent of physicians' malpractice liability in each state, where we assume that $L_3 > 0$, indicating that as k increases, physicians incur greater legal liability, e.g., because states impose fewer limitations on physician liability. We also assume that $L_{13} \leq 0$, indicating that increasing the extent of legal liability in the environment either makes the effect of revenue-generating effort on legal liability stronger (i.e., more negative) or leaves it unchanged. Likewise, we assume that $L_{23} \geq 0$, indicating that increasing the extent of legal liability in the environment either makes the effect of cost control effort on legal liability stronger (i.e., more positive) or leaves it unchanged.

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[9] Capitation contracts impose significant risk on a physician or physician practice because under a capitation contract the MCO pays the practice or the physician a fixed amount per enrolled member per month, and the practice or physician bears the uncertainty with respect to the actual cost of care. That is, if a physician paid on a capitated basis incurs costs in patient care greater than the contractually agreed capitation payment, then the physician must absorb the excess cost. This holds whether the additional cost arises from an unfavorable state (unusually sick patients) or because the physician consumes excessive resources in patient care. At the other extreme, under fee-for-service (FFS) arrangements physicians are paid for specific services rendered, and therefore bear less of the risk of the actual cost of care. Physicians paid under FFS do face revenue uncertainty, but as noted earlier, we assume this risk is small relative to the risk related to cost and malpractice liability. Under an FFS arrangement, if actual expenditures are larger than expected because the physician performs additional tests and procedures, then the MCO's payment to the physician will generally cover the additional cost because the reimbursement is based on established fees for each service rendered.

[10] See Feltham and Xie (1994). The liability function needs to be separable and also linear or convex in the actions to ensure that the sum $L\,(.) + C\,(.)$ is convex.

Next, let the error term $e_L \sim N(0, \text{var}_L)$ represent the effect of all other factors on the extent of the physician's liability loss, where we assume that the revenue and liability error terms are uncorrelated, i.e., $\text{cov}(e_1,e_L) = 0$. Finally, to reflect the fact that more complex cases are likely to result in unfavorable (low) cost control outcomes and high legal liability outcomes, we assume a negative relation between the error terms e_2 and e_L, as $\text{cov}(e_2,e_L) < 0$. This, in turn, implies $-1 < \rho_{2L} < 0$, where ρ_{2L} is the correlation between error terms e_2 and e_L.[12]

Expressing the agent's expected utility in terms of the certainty equivalent, and denoting U^0 as the physician's reservation expected utility, the risk-neutral MCO's problem in contracting with the physician then becomes:[13]

$$\underset{\alpha,\beta_1,\beta_2,a_1,a_2}{Max} \quad E[x(a_1,a_2) - w(y_1,y_2)]$$

subject to:

$$E[w(y_1,y_2) - L(a_1,a_2,k)] - .5r\,\text{var}[w(y_1,y_2) - L(a_1,a_2,k)] - C(a_1,a_2) \geq U^0 \tag{1}$$

$$a_1,a_2 \in \arg\max E[w(y_1,y_2) - L(a_1,a_2,k)] - .5r\,\text{var}[w(y_1,y_2) - L(a_1,a_2,k)] - C(a_1,a_2) \tag{2}$$

where Equation (1) is the physician's reservation expected utility constraint and Equation (2) ensures incentive compatibility.

Assume without loss of generality that $U^0 = 0$. Let the fixed payment α be chosen to satisfy Equation (1) as an equality, and substitute from Equation (1) for $E[w(y_1,y_2)]$ into the objective function. Writing the MCO's expected payoff (x) in terms of actions and the corresponding sensitivities ($x = f_1 a_1 + f_2 a_2$) and expanding the expression for $.5r \operatorname{var}[w(y_1,y_2)]$, the MCO's problem becomes:

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[11] Kessler and McClellan (1996, 2000).

[12] U.S. Congress, Office of Technology Assessment (OTA 1993) reports that the probability of litigation is increasing in the severity and cost of a patient's medical injury.

[13] According to this model specification, legal liability enters the MCO's objective function only indirectly via its role in the physician's individual rationality constraint. This approach reflects the fact that the Employment Retirement Income Security Act of 1974 (ERISA) exempts HMOs from lawsuits based on negligent patient care. However, this protection from liability has been under attack in recent years through various efforts designed to increase HMOs' legal liability (Butler 2001). An alternative specification of the MCO's objective function would include a term to capture the MCO's potential legal liability for negligent patient care. Incorporating such a term would increase the MCO's returns from the physician's revenue generating effort and decrease the returns to cost control effort, which would increase the ratio of optimal actions, a_1/a_2, and in turn, reduce the ratio of optimal incentive weights, β_2/β_1. Assuming that the new effect did not interact with other components of the MCO's expected return, the results in Propositions 1 and 2 would be stronger but directionally the same.

$$\begin{aligned}\underset{\beta_1,\beta_2,a_1,a_2}{Max}\ & f_1 a_1 + f_2 a_2 - C(a_1,a_2) - L(a_1,a_2,k) - .5r[\beta_1^2 \operatorname{var}_1 + \beta_2^2 \operatorname{var}_2 \\ & +2\beta_1\beta_2 \operatorname{cov}(e_1,e_2) - 2\beta_1 \operatorname{cov}(e_1,e_L) - 2\beta_2 \operatorname{cov}(e_2,e_L) + \operatorname{var}_L].\end{aligned}$$

Using the assumptions that $g_{12} = g_{21} = \operatorname{cov}(e_1,e_2) = \operatorname{cov}(e_2,e_L) = 0$, the physician's incentive compatibility constraint for action a_1 becomes:

$$a_1 \in \arg\max E[\beta_1(g_{11}a_1) + \beta_2(g_{22}a_2) - (.5a_1^2 + .5a_2^2) - L(a_1,a_2,k)],$$

where the variance terms drop out because they are not affected by the physician's actions. The expression for a_2 is similar. The physician's first-order conditions with respect to a_1 and a_2 become:

$$\beta_1 g_{11} - a_1 - L_1\,(a_1,a_2,k) = 0; \tag{3}$$

$$\beta_2 g_{22} - a_2 - L_2\,(a_1,a_2,k) = 0. \tag{4}$$

For a given contract, from Equations (3) and (4), the earlier assumptions that $L_{13} \leq 0$ and $L_{23} \geq 0$ imply that an increase in legal liability (k) would cause a weak increase in revenue generation effort (a_1) and a weak decrease in cost control effort (a_2).

Next, substitute into the objective function for a_1 and a_2 from Equations (3) and (4) to give the MCO's objective function without constraints:

$$\begin{aligned}\underset{\beta_1,\beta_2}{Max}\ & f_1[\beta_1 g_{11} - L_1] + f_2[\beta_2 g_{22} - L_2] - .5[\beta_1 g_{11} - L_1]^2 - .5[\beta_2 g_{22} - L_2]^2 \\ & - L[(\beta_1 g_{11} - L_1), (\beta_2 g_{22} - L_2), k] - .5r[\beta_1^2 \operatorname{var}_1 + \beta_2^2 \operatorname{var}_2 - 2\beta_2 \operatorname{cov}(e_2,e_L) \\ & + \operatorname{var}_L].\end{aligned} \tag{5}$$

We use this model to analyze how the optimal weights (β_1,β_2) in the MCO-physician contract vary across different MCO-physician environments. First, in Proposition 1, we examine how the ratio of the incentive weights (β_2/β_1) varies with the noisiness of the cost measure as reflected by var_2. Second, in Proposition 2, we analyze how the ratio changes with the magnitude of the association between the cost control and legal liability outcomes, as reflected by $\operatorname{cov}(e_2,e_L)$.

Proposition 1: As the cost measure becomes noisier (i.e., as var_2 increases), the MCO will decrease β_2/β_1.

Proposition 1 (proofs are in Appendix B) predicts that as the cost control measure becomes noisier, the MCO will place less weight on the cost control measure (less capitation) relative to the weight on the revenue-generating measure, which means that β_2/β_1 will decline. The intuition for Proposition 1 comes from the precision effect (Banker and Datar 1989), whereby less precise (greater variance) performance measures receive less weight. Datar et al.'s (2001) Proposition 3 also isolates this effect, but with the assumption that the two performance measures are perfectly congruent with each other. In contrast, Proposition 1 above holds independent of the level of congruence between the performance measures. Proposition 1 leads to the following hypothesis.

H1: MCOs will rely more on capitation in contracts with PCPs than with specialists.

Hypothesis 1 follows from Proposition 1 together with the assumption that specialists face greater cost uncertainty than PCPs, as we now describe. Most physician-patient encounters involve relatively simple medical conditions, and the practice of medicine that deals with these events is commonly described as general medicine or primary care (Phelps 1997). Primary care physicians (PCPs) practice family/general medicine, internal medicine, or pediatrics, and act as gatekeepers in some managed care systems.[14] PCPs provide an initial diagnosis, initiate treatment in less complex cases, and ensure overall continuity of care, often seeing patients several times.

Specialists are physicians with expertise in particular aspects of medical care. Specialists provide consultation services to primary care physicians in managing more complex cases. They also treat patients who are not responding to initial primary care-based treatment, and provide ongoing specialty care for the most severe or complicated cases. Specialists generally provide more complex treatments and medical procedures, and are more likely to see patients who are sicker and require more complex interventions (Meyer et al. 2000; Katon et al. 2001). Specialists likewise face a greater risk of adverse outcomes from their medical treatment than do PCPs. For example, the Harvard Medical Malpractice Study (1990) found that the rate of adverse events (injury attributable to medical care) varied from 3.6 percent for general medicine (primary care) to 7.0 percent for general surgery, 9.9 percent for neurosurgery, 10.8 percent for thoracic surgery, and 16.1 percent for cardiac and vascular surgery. Given that specialists face greater task-related uncertainty than PCPs, it follows from Proposition 1 that this difference in task characteristics should lead MCOs to rely more on capitation in contracts with PCPs than in contracts with specialists.

Legal Liability Risk and Optimal Incentive Weights

Another influence on the form of MCO-physician contracts in our model is the nature and extent of physicians' legal liability. Cross-sectional differences in physicians' malpractice liability could influence the optimal incentive weights (β_2/β_1) in two ways. First, an increase in the legal liability parameter, k, would represent greater malpractice exposure for physicians. However, from Appendix B it can be seen that varying the level of legal liability alone should have no effect on the optimal incentive weights, (β_2/β_1).[15] Second, an increase in legal liability could also affect the optimal incentive weights through what we term the *covariance effect* of legal liability. The covariance effect of legal liability is represented by the $\text{cov}(e_2,e_L)$ term in the MCO's objective function in Expression (5). This $\text{cov}(e_2,e_L)$ term measures the strength of the association, which we earlier assumed to be negative ($\text{cov}(e_2,e_L) < 0$), between the cost control error term (e_2) and the legal liability error term (e_L). Proposition 2 relates the magnitude of the covariance effect of legal liability, as reflected in the correlation ρ_{2L} between cost-control and legal liability outcomes, to the optimal incentive weights, β_2/β_1.

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[14] Ob/Gyn physicians are generally classified as specialists even though they share some characteristics with PCPs, e.g., members may have direct access to Ob/Gyn physicians (Kongstvedt 2001, 110).

[15] From the physicians' first-order conditions, Equations (3) and (4), an increase in the extent of malpractice liability (k) with all else held constant would produce a weak increase in the physician revenue-generating effort and a weak decrease in cost-control effort. However, from Equation (8) in Appendix B, a change in k will not affect the MCO's optimal choice of incentive weights, (β_2/β_1). The reason is that the physician's effort adjustments produce equal and offsetting effects on the MCO's cost of inducing physician effort and, hence, the optimal incentive weights would remain unchanged.

Proposition 2: As the association between unfavorable cost control outcomes and malpractice liability realizations becomes stronger (i.e., as ρ_{2L} becomes more negative), MCOs will reduce β_2/β_1, reflecting less reliance on cost control (capitation) in physicians' contracts.

Proposition 2 predicts that as the association between unfavorable cost control outcomes (reflecting adverse patient outcomes) and the extent of malpractice liability becomes stronger, MCOs will rely less on capitation. The assumption that $\rho_{2L} < 0$ is based on evidence that the probability of litigation is increasing in the severity and costliness of a patient's medical injury (OTA 1993). Further, we expect the absolute value of this negative association between the extent of malpractice liability and cost control outcomes to be strongest for specialists practicing in states with greater legal liability. That is, when specialists experience unfavorable cost control outcomes (high cost of patient care) in high legal liability states, they are also likely to incur a large malpractice liability loss, L. We expect this association to be weaker for PCPs and for specialists practicing in states with less legal liability. That is, for specialists practicing in lower legal liability states, unfavorable cost control outcomes are less likely to also produce high legal liability losses because the probability of litigation and the expected loss from litigation will be lower in low liability states. Finally, for PCPs, large legal liability losses are less likely to be associated with unfavorable cost control outcomes because such cases are very likely to be transferred to the care of specialists and, hence, the PCP will not bear the increased cost. Hypothesis 2 formalizes this prediction:

H2: MCOs will employ less capitation in physician contracts for specialists in states with greater legal liability than in contracts with PCPs or in contracts with specialists in states with less legal liability.

Hypothesis 1 predicted that noisier cost control measures for specialists than for PCPs would result in less capitation for specialists, independent of the legal liability environment. Hypothesis 2 analyzes how the legal liability environment interacts with the extent of noise in cost control measures. Specifically, because the extent of capitation should be least when the absolute value of ρ_{2L} is greatest, i.e., when the association between unfavorable cost control outcomes and legal liability outcomes is strongest, H2 predicts the least capitation for specialists in states with high physician malpractice liability.

III. DATA, VARIABLES, AND MODEL

Individual Physician Specialties and Practice Compensation Arrangements

We obtained our basic sample of physicians and their practices' financial incentive data from a national telephone survey of physicians, administered for the Center for Studying Health System Change's Community Tracking Study by the Gallup Organization in 1996–97 and repeated with similar national physician samples in 1998–99 and 2000–01 (Center for Studying Health System Change 2002). Gallup selected random samples of PCPs and specialists for each of 60 sites (51 metropolitan areas and 9 non-metropolitan areas) across 33 states. The three surveys yielded 37,238 responses from full-time physicians.

The surveys classify physicians into seven care categories: family/general, internal medicine, pediatrics, medical specialties, surgical specialties, Ob/Gyn, and psychiatry. We excluded psychiatrists because they do not perform direct medical procedures, as well as physicians employed by universities, government agencies, hospitals and staff, or group model HMOs because they are likely to face different incentives than those in a typical physician practice. We also dropped physicians in multi-specialty practices so that we could unambiguously characterize each physician's practice in terms of the type of tasks performed. We further dropped physicians who did not respond to the question about their practice size, yielding a final sample of 18,680 survey responses.[16]

Malpractice Cost

The Malpractice Geographic Practice Cost Index (MGPCI) is the Medicare index of physician malpractice insurance costs, and reflects the relative cost of physicians' malpractice insurance across our sample sites (Zuckerman and Norton 1994; Health Care Financing Administration [HCFA] 2001, HCFA-1120-P 2001).

State Malpractice Laws

We collected data on each state's malpractice provisions from a 1993 report by the OTA and updated this information to 1995 (McCullough, Campbell & Lane 2003).

We used these data to develop the following measures.

(1) Percentage of capitation revenue (*PCAPREV*): *PCAPREV* is our primary dependent measure, reflecting the extent to which the practice bears the residual cost risk of patient care. Each physician reported the proportion of total revenue *for the entire physician practice* received as capitation payments.[17] The survey question asked:

> Now, thinking again about the patient care revenue from ALL sources received by the practice in which you work, what percentage is paid on a capitated or other prepaid basis?

(2) Physician Specialty (*PHY-DUMMY*): We created dummy variables for family/general (*FAM/GEN*), internal medicine (*INTERNAL*), pediatrics (*PED*), medical specialists (*MEDSPEC*), and surgical specialists (*SUGSPEC*). *OB/GYN* is the omitted physician specialty.

(3) Specialist and High Malpractice Liability (*SPEC-DUMMY* × *HIMGPCI*): MGPCI is the 1998 Malpractice Geographic Practice Cost Index, reflecting the extent of legal liability in each location. We created three indicator variables, one for each of the three physician specialist categories, in combination with high legal liability defined as above median values of MGPCI. For example, *MEDSPEC* × *HIMGPCI* = 1 for medical specialists practicing in states with above median values of MGPCI.

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[16] Although a different random sample was chosen for each of the three surveys, some physicians were selected more than once. We estimate that of our final sample of 18,680 physician responses, 8,214 responses came from 2,738 physicians who each participated in all three surveys, 4,540 responses came from 2,270 physicians who each participated in two surveys, and 5,926 responses came from physicians who participated in only one survey. The surveys collect data from individual physicians who provide information about financial arrangements for their physician practice, so our data is likely to include at least some instances of multiple physician responses from the same physician practice in a given year, as well as many instances of the same physician responding in more than one year. Therefore, the total number of physician responses overstates the number of independent observations in our sample. However, consistent with 51.9 percent of our observations coming from practices comprised of only one or two physicians, subsequent robustness checks confirm that this effect does not significantly influence our results.

[17] We use the log transformation, $\ln(1 + PCAPREV)$ because the distribution of capitation revenue is skewed to the right.

(4) State Malpractice Provisions: Following OTA (1993) and Kessler and McClellan (1996, 2000), we create two alternative measures of the extent of physicians' malpractice liability by state. "Direct caps" on malpractice liability limit malpractice damage awards and the potential to collect from multiple parties for the same damages. The more of these direct caps on liability apply in a particular state, the smaller is *DIRLIAB*, the measure of physicians' malpractice liability in that state. Specifically, we define *DIRLIAB* as equal to 2 when no such limits (caps) are in effect, 1 when one cap applies, and 0 in states that impose both forms of direct caps on physicians' malpractice liability. "Indirect caps" on malpractice liability require mandatory periodic payments to plaintiffs rather than a lump sum payment, constrain joint liability, and limit attorneys' fees.

We define *INDLIAB* similarly to *DIRLIAB*, with values of 3, 2, 1, and 0, depending on how many indirect limitations on physicians' malpractice liability are effective in each state. In the regression model, we use an indicator variable *HIGH-DIRLIAB* which equals 1 for above median *DIRLIAB*, and otherwise 0. In a similar manner, we use an indicator variable *HIGH-INDLIAB* for *INDLIAB*.

(5) Control variables: We include controls for the following factors that may also influence the extent of cost risk imposed on physician practices.
  (a) Percentage of managed care revenue (*PMC*): *PMC* reflects the extent of the practice's involvement in managed care, and is measured as the proportion of the practice's total revenue received from managed care organizations. *PMC* proxies for the bargaining power of a physician practice relative to MCOs because as *PMC* increases, the physician practice becomes more dependent on managed care.
  (b) Practice size (*SIZE*): log (*SIZE*) is the natural log of the number of physicians in the practice. Leone (2002) argues that risk premiums will be smaller for larger practices because they are better able to share the financial risk from capitation contracts. Ittner et al. (2006) propose that smaller practices can monitor individual physicians more effectively than larger practices. Both arguments suggest that larger practices should rely more on capitation.
  (c) Effect of practice guidelines (*EFGUID*): MCOs frequently use formal, written guidelines for the delivery of patient care in specific circumstances as part of their control systems. For example, the guidelines may specify when certain tests should be performed. Physicians rated the extent to which such guidelines affected their practice of medicine on a scale from 0 (no effect) to 5 (very large effect).
  (d) Survey Site Population (*POPULATN*): *POPULATN* is the population for the survey site (U.S. Census Bureau 2000). A larger population in the local market facilitates capitation arrangements, which typically require at least a minimum number of patients.

We pooled observations from the three surveys and conducted cross-sectional data analyses using the following OLS regression model:[18]

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[18] We also include dummies for observations from the second and third surveys.

$$\ln(1 + PCAPREV_i) = \alpha + \sum_{j=1}^{5} \delta_j PHY\text{-}DUMMY_{ij} + \sum_{k=1}^{3} \gamma_k (SPEC\text{-}DUMMY_{ik} \times HIMGPCI_i) + \sum \text{Control Variables}$$

where:

$i = 1,\ldots,18{,}680$ refers to the practice in which the surveyed physician works;
$j = 1,\ldots,5$ refers to the type of PCP or specialist for each practice; and
$k = 1,2,3$ refers to the type of specialist.

V. RESULTS

Descriptive Statistics

The *PMC* (percentage of managed care revenue) column in Table 1 shows that for each of the six physician categories, physician practices receive a significant percentage (mean 35–59 percent, median 30–60 percent) of their total revenue from MCOs. From the *PCAPREV* column of Table 1, PCP practices (*FAM/GEN*, *INTERNAL*, *PED*) receive on average 19–23 percent (median 10–15 percent) of their revenue in the form of capitation payments versus only 6–9 percent (median 0 percent) for specialists (*OB/GYN*, *MEDSPEC*, *SUGSPEC*). The next two columns describe the size of all physician practices (*SIZE*), as well as the size of practices after dropping 291 practices with over 100 physicians (*SIZE* (*w*)). Based on the *SIZE* (*w*) measure, most practices are small with an overall mean (median) of 5.81 (2.00) physicians per practice.[19] The last column of Table 1 shows that for each physician category, respondents on average perceive that practice guidelines exert a small (2) to moderate (3) effect on their practice of medicine, with the largest perceived effect in Ob/Gyn.

The stratified national sample includes responses from 60 survey locations. With respect to our primary dependent variable, percentage of capitation revenue (*PCAPREV*), the sample locations (tables not reported) had a mean value of 14.93 percent and varied from a low of 1.49 percent in West Central Alabama to a high of 33.78 percent in Modesto, California. The physician malpractice insurance premium index (MGPCI) also displays considerable variation around the mean of 0.968 with a low of 0.399 in Central Arkansas and a high of 3.002 in Detroit, Michigan. The measures of physicians' malpractice liability based on direct (*DIRLIAB*) and indirect (*INDLIAB*) limits on liability also reflect significant variation around their mean values of 1.23 and 1.72, respectively.

Table 2 reports the Pearson correlations among our variables. The two measures of physicians' liability based on specific liability limits (*DIRLIAB*, *INDLIAB*) display a strong, positive correlation (0.56). The correlations of various variables with the survey site population (*POPULATN*) indicate that larger sites tend to have more managed care (*PMC*), more capitation (*PCAPREV*), higher legal liability costs (*MGPCI*), but lower legal liability when measured by the number of direct (*DIRLIAB*) and indirect (*INDLIAB*) reforms that reduce liability. Related to this final observation, the negative correlations of −0.27 and −0.37, respectively, between the malpractice liability cost measure (*MGPCI*) and the malpractice liability measures based on specific reforms (*DIRLIAB*, *INDLIAB*) are unexpected.

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[19] For approximately 84.7 percent of our observations, the physicians jointly own the practice. The remaining 15.3 percent of all observations include approximately 11.7 percent of all cases in which physicians share ownership with some outside entity, and 3.6 percent of all cases in which the practice is owned entirely by an outside entity. Hence, *SIZE* (number of physicians in the practice) reflects both practice size and also ownership structure (number of owners) for most of our observations.

TABLE 1
Physician and Practice-Level Descriptive Statistics for Surveys 1–3[a]

Physician's Specialty	Descriptive Statistics	*PMC* (%)	*PCAPREV* (%)	*SIZE*	*SIZE*(*w*)	*EFGUID*
FAM/GEN	Mean	44.54	19.89	9.16	5.16	2.39
	Median	45.00	10.00	2.00	2.00	3.00
	25th Percentile	25.00	0.00	1.00	1.00	2.00
	75th Percentile	65.00	30.00	4.00	4.00	3.00
	Std. Deviation	26.33	23.53	44.88	11.44	1.27
	n	4,875	4,875	4,875	4,802	4,875
INTERNAL	Mean	40.77	18.92	13.01	6.25	2.34
	Median	40.00	10.00	2.00	2.00	2.00
	25th Percentile	20.00	0.00	1.00	1.00	1.00
	75th Percentile	60.00	30.00	5.00	5.00	3.00
	Std. Deviation	25.44	22.71	60.14	13.56	1.30
	n	3,425	3,425	3,425	3,352	3,425
PED	Mean	59.40	22.91	9.53	5.90	2.45
	Median	60.00	15.00	3.00	3.00	3.00
	25th Percentile	40.00	0.00	1.00	1.00	2.00
	75th Percentile	80.00	40.00	6.00	6.00	3.00
	Std. Deviation	23.95	23.81	38.98	10.63	1.28
	n	2,664	2,664	2,664	2,622	2,664
OB/GYN	Mean	53.59	8.98	5.73	4.27	2.69
	Median	50.00	0.00	2.00	2.00	3.00
	25th Percentile	40.00	0.00	1.00	1.00	2.00
	75th Percentile	75.00	5.00	5.00	4.50	4.00
	Std. Deviation	24.76	19.56	29.60	7.80	1.55
	n	881	881	881	877	881
MEDSPEC	Mean	35.22	8.27	14.67	7.93	2.27
	Median	30.00	0.00	4.00	3.00	2.00
	25th Percentile	20.00	0.00	1.00	1.00	1.00
	75th Percentile	50.00	10.00	9.00	8.00	3.00
	Std. Deviation	22.81	15.89	55.82	13.25	1.28
	n	3,467	3,467	3,467	3,385	3,467
SUGSPEC	Mean	37.34	5.53	6.48	4.50	2.25
	Median	35.00	0.00	2.00	2.00	2.00
	25th Percentile	20.00	0.00	1.00	1.00	1.00
	75th Percentile	50.00	3.00	5.00	5.00	3.00
	Std. Deviation	23.18	14.30	34.41	7.83	1.40
	n	3,368	3,368	3,368	3,351	3,368
Total	Mean	43.37	14.87	10.30	5.81	2.36
	Median	40.00	4.00	2.00	2.00	2.00
	25th Percentile	25.00	0.00	1.00	1.00	1.00
	75th Percentile	60.00	25.00	5.00	5.00	3.00
	Std. Deviation	25.81	21.59	47.40	11.47	1.31
	n	18,680	18,680	18,680	18,389	18,680

(*continued on next page*)

TABLE 1 (Continued)

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[a] The 1996–1997 survey, the 1998–1999 survey, and the 2000–2001 survey collected data as of 1995, 1997, and 1999, respectively.

FAM/GEN = family/general practice;
INTERNAL = internal medicine;
PED = pediatrics;
OB/GYN = obstetrics and gynecology;
MEDSPEC = medical specialties;
SUGSPEC = surgical specialties;
PMC = the percentage of a practice's total patient care revenue that is received from all managed care organizations (MCOs);
PCAPREV = the percentage of a practice's total patient care revenue received on capitated or other prepaid basis;
SIZE = the number of physicians working for the practice;
SIZE(w) = the number of physicians working for the practice after dropping practices with more than 100 physicians;
EFGUID = the extent to which formal/written practice guidelines from HMOs or medical groups affect the physicians' practice (0 = no effect, 1 = very small, 2 = small, 3 = moderate, 4 = large, and 5 = very large); and
n = the number of physicians in the sample.

TABLE 2
Pearson Correlation Matrix for Surveys 1–3[a]
(n = 18,680)

	PMC	*PCAPREV*	*MGPCI*	*DIRLIAB*	*INDLIAB*	*SIZE*	*EFGUID*
PMC	1						
PCAPREV	0.47**	1					
MGPCI	0.03**	0.05**	1				
DIRLIAB	−0.07**	−0.03**	−0.27**	1			
INDLIAB	−0.06**	−0.08**	−0.37**	0.56**	1		
SIZE	0.05**	0.10**	−0.03**	0.01	0.01	1	
EFGUID	0.13**	0.14**	0.04**	−0.03**	−0.03**	0.00	1
POPULATN	0.10**	0.21**	0.09**	−0.16**	−0.21**	0.02*	0.04**

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*, ** Correlation is significant at the 0.01 and 0.001 levels, respectively (two-tailed).

[a] The 1996–1997 survey, the 1998–1999 survey, and the 2000–2001 survey collected data as of 1995, 1997, and 1999, respectively.

PMC = the percentage of a practice's total patient care revenue received from all managed care organizations;
PCAPREV = the percentage of a practice's patient care revenue paid on a capitation or other prepaid basis;
MGPCI = the Malpractice Geographical Practice Cost Index used by the Medicare for cost reimbursement to physicians;
DIRLIAB = malpractice liability as reduced by reform laws that *directly* limit the amount of damage awards to patient(s); Larger values indicate greater legal liability;
INDLIAB = malpractice liability as reduced by reform laws that *indirectly* limit the frequency of malpractice litigation; Larger values indicate greater legal liability;
SIZE = the number of physicians working for the practice;
EFGUID = the extent to which formal/written practice guidelines from HMOs or medical groups affect the physician's practice (0 = no effect, 1 = very small, 2 = small, 3 = moderate, 4 = large, and 5 = very large);
POPULATN = the population of the Metropolitan Statistical Area (MSA) in which the physician practice is located; and
n = the number of physicians in the sample.

They indicate that malpractice liability costs are greater in states that have experienced more extensive malpractice reforms, which translate to smaller values of *DIRLIAB* and *INDLIAB*. One potential interpretation is that states with higher liability costs are more likely to impose caps on malpractice litigation. Finally, Table 3 shows that the percentage of capitation revenue is positively associated with the perceived effect of practice guidelines.

OLS Regression Results

In OLS Models 1–3 in Table 4, the four control variables with hypothesized signs are all statistically significant ($p < 0.001$) with the predicted signs in all models. In particular, MCO-physician practice contracts rely more on capitation as the extent of managed care (*PMC*) in the local market increases, as the physician practice becomes larger (*SIZE*), when practice guidelines (*EFGUID*) exert a greater effect on physicians, and as the size of the market (*POPULATN*) increases. The final two control variables without hypothesized signs are dummy variables for the second (*S2 DUMMY*) and third (*S3 DUMMY*) survey administrations. Both are negative, consistent with a decline in the overall reliance on capitation over the sample period.

Hypothesis 1

Hypothesis 1 predicts that MCOs will use capitation more extensively in contracts with PCPs than in contracts with specialists. Consistent with H1, in OLS Model 1 in Table 4 the estimated coefficients on the dummy variables for PCPs (*FAM/GEN, INTERNAL, PED*) are significantly larger ($p < 0.001$) than the coefficients for specialists (*MEDSPEC, SUGSPEC*), indicating that specialists bear less cost risk (less capitation) in MCO-physician contracts than PCPs. More specifically, the notes below Table 4 report the results of F-tests comparing the magnitude of the estimated coefficients of each of the three PCP indicator variables to the magnitude of the estimated coefficients of each of the two specialist indicator variables. The results indicate that each estimated PCP coefficient is larger than each estimated specialist coefficient at the $p < 0.001$ level. The results for Models 2 and 3 in Table 4 are very similar to those for Model 1. Given the previously cited evidence that

TABLE 3
Mean Percentage of Capitation Revenue by Perceived Effect of Practice Guidelines for Surveys 1–3[a]

Effect of Practice Guidelines (*EFGUID*)	n	Percentage of Revenue from Capitation (*PCAPREV*)
No Effect (0)	2,089	8.77
Very Small (1)	2,624	11.59
Small (2)	4,798	13.97
Moderate (3)	5,717	16.57
Large (4)	2,657	19.40
Very Large (5)	795	19.73
Total	18,680	14.87

[a] The 1996–1997 survey, the 1998–1999 survey, and the 2000–2001 survey collected data as of 1995, 1997, and 1999, respectively.

Based on physicians' responses, *EFGUID* measures the extent to which physicians perceive that formal written guidelines affect how they practice medicine. *PCAPREV(%)* is the percentage of a practice's total patient care revenue received on a capitated or other prepaid basis.

TABLE 4
Effect of Physician Specialty and Malpractice Costs (MGPCI) on Physician Practice's Capitation Revenue for Surveys 1–3[a] (OLS Estimates)

Dependent variable: log (1 + *PCAPREV*)

Independent Variables	**Model 1 Coefficient (t^b)**		**Model 2 Coefficient (t^b)**		**Model 3 Coefficient (t^b)**	
Intercept	−3.14	−26.23**	−3.03	−22.71**	−3.14	−26.36**
FAM/GEN (δ_1)	1.32	24.32**	1.13	14.26**	1.31	24.27**
INTERNAL (δ_2)	1.32	23.67**	1.13	14.07**	1.32	23.71**
PED (δ_3)	1.21	20.73**	1.02	12.40**	1.21	20.75**
MEDSPEC (δ_4)	0.44	7.81**	0.33	3.99**	0.44	7.81**
SUGSPEC (δ_5)	0.15	2.65*	0.09	1.09	0.14	2.61*
PMC	0.02	44.76**	0.02	44.87**	0.02	44.78**
log (*SIZE*)	0.18	21.55**	0.18	21.20**	0.18	21.23**
EFGUID	0.12	15.80**	0.12	15.80**	0.12	15.82**
log (*POPULATN*)	0.19	24.25**	0.19	24.89**	0.19	24.46**
S2 DUMMY	−0.05	−2.03	−0.05	−2.13	−0.05	−2.05
S3 DUMMY	−0.34	−13.86**	−0.33	−13.71**	−0.34	−13.75**
HIMGPCI					−0.05	−2.31
*MEDSPEC*HIMGPCI* (γ_1)			−0.16	−3.55**		
*SUGSPEC*HIMGPCI* (γ_2)			−0.27	−6.47**		
*OB/GYN*HIMGPCI* (γ_3)			−0.34	−3.42**		
Adjusted R^2	.30		.31		.30	
n	18,680		18,680		18,680	

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*, **Indicate two-tailed significance at the 0.01 and 0.001 levels, respectively.

[a] The 1996–1997 survey, the 1998–1999 survey, and the 2000–2001 survey collected data as of 1995, 1997 and 1999, respectively.

[b] We report asymptotic t-statistics based on White (1980) standard errors.

F-test for H1 (*PCP* > Specialists) in Model 1:

H_0: $\delta_1 = \delta_4$ 800.58 ($p < 0.001$) H_0: $\delta_1 = \delta_5$ 1445.45 ($p < 0.001$)
H_0: $\delta_2 = \delta_4$ 699.67 ($p < 0.001$) H_0: $\delta_2 = \delta_5$ 1238.35 ($p < 0.001$)
H_0: $\delta_3 = \delta_4$ 438.25 ($p < 0.001$) H_0: $\delta_3 = \delta_5$ 839.81 ($p < 0.001$)

Dummy variables for three types of primary care practices are:
FAM/GEN (Family/General), *INTERNAL* (Internal Medicine), and PED (Pediatrics).
Dummy variables for specialist practices are:
MEDSPEC (Medical specialties), *SUGSPEC* (Surgical specialties), and OB/GYN (obstetrics and gynecology).

PCAPREV = the percentage of a physician practice's total patient care revenue paid on capitated or other prepaid basis;
MGPCI = the Malpractice Geographical Practice Cost Index used for Medicare cost reimbursement;
HIMGPCI = a dummy variable for high MGPCI, coded as 1 for the above median MGPCI, and 0 for the below median MGPCI;
PMC = the percentage of a practice's total patient care revenue received from all managed care organizations (MCOs);
SIZE = the number of physicians working for the practice;
EFGUID = the extent to which formal/written MCO practice guidelines affect the physician's practice (0 = no effect, 1 = very small, 2 = small, 3 = moderate, 4 = large, and 5 = very large);
POPULATN = population of the Metropolitan Statistical Area (MSA) in which the physician practice is located;
S2 DUMMY and *S3 DUMMY* = indicator variables for Surveys 2 (1998–1999) and Survey 3 (2000–2001), respectively; and
log = natural log.

PCPs face less exogenous cost risk than specialists, the results in Table 4 confirm the prediction in H1 that MCOs will rely less on capitation for specialists because cost measures are noisier for specialists than for PCPs.

Hypothesis 2

We test H2 using Model 2 in Table 4. Model 2 adds the three specialist-high legal liability categorical variables (*MEDSPEC*HIMGPCI*, *SUGSPEC*HIMGPCI*, *OB/GYN* HIMGPCI*) to Model 1 to test the effect of a high negative correlation of cost control outcomes and liability outcomes. For example, *MEDSPEC*HIMGPCI* = 1 for medical specialists practicing in states with above the median legal liability, i.e., in environments in which we expect the strongest association of cost control and legal liability outcomes, and *MEDSPEC*HIMGPCI* = 0 otherwise, where we expect a weaker association of cost control and legal liability.

Hypothesis 2 predicts that MCOs will employ less capitation for specialists practicing in states with high legal liability than for PCPs or for specialists practicing in states with low legal liability. Hence, H2 implies that the three specialist-high legal liability categorical variables will have negative coefficients. The results in Model 2 of Table 4 show that, consistent with H2, the coefficients on each of the three specialist-high legal liability categorical variables (*MEDSPEC*HIMGPCI*, *SUGSPEC*HIMGPCI*, *OB/GYN*HIMGPCI*) are negative and statistically significantly different from zero at the $p < 0.001$ level.

In Model 3 of Table 4 we replace the three specialist-high legal liability categorical variables with a single high legal liability categorical variable (*HIMGPCI*). The estimated coefficient is negative but not statistically significantly different from zero at the 0.01 level, consistent with the notion that the combination of specialists and high legal liability is a better proxy for environments with strong correlations of unfavorable cost control outcomes and high legal liability than is high legal liability alone. This result is also consistent with the implication of our model that changing legal liability alone (*MGPCI*) should not change the extent of capitation (see footnote 15).[20]

The preceding results are consistent with the prediction in Proposition 2 that as the negative association between cost control and malpractice liability becomes stronger, MCOs will rely less on capitation. Such a stronger association characterizes the environments facing specialists practicing in states with few limitations on the extent of physician malpractice liability, i.e., in states with high legal liability for physicians.

To illustrate the economic magnitude of the preceding estimated effects, we compare the estimated percentage capitation for a Family/General PCP versus a Surgical Specialist practicing in a low versus a high legal liability environment. Fixing all independent variables except categorical variables at their median values, Model 2 of Table 4 yields estimated capitation percentages of 8.3 percent for Family/General physicians versus 2.0 percent for Surgical Specialists in a low legal liability environment and 1.4 percent for Surgical Specialists in a high legal liability environment.

Finally, although capitation provides desirable cost control incentives, high levels of capitation can also create potential legal and political exposure for MCOs if physicians reduce services to the point that the quality of care is significantly compromised. One potential response by MCOs is to impose practice care guidelines to minimize problems with physicians providing too little care, and a second response is to incorporate quality clauses in physician contracts (Evans et al. 2005). The significant positive coefficients for

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[20] The results for both hypotheses are similar if *HIMGPCI* is replaced in Model 3 of Table 4 with the continuous variable *MGPCI*.

EFGUID in Table 4 are consistent with the results in Table 3, reflecting a positive association between the extent of capitation and physicians' perceptions of the effect of guidelines on their practice decisions. This association between capitation and the effect of guidelines is consistent with MCOs combining high-powered incentives (greater capitation) with less delegation of discretion to physicians as the MCOs rely more on practice guidelines.

Robustness Tests

First, we checked the effect of measuring high legal liability as a dichotomous variable (*HIMGPCI*) by replacing the three specialist-high *MGPCI* categorical variables in Model 2 of Table 4 with three specialist-*MGPCI* interactions, and replacing *HIMGPCI* in Model 3 of Table 4 with *MGPCI.* These alternative specifications yield qualitatively the same results. Specifically, the coefficients for the three specialist-*MGPCI* interactions are negative and statistically significant at $p < 0.001$, whereas the mean effect of *MGPCI* is positive but not significantly different from zero at the 0.01 level.

Second, because 8,549 of our 18,680 total observations have 0 values for the percentage of capitation, creating a censored data set, we reran Model 2 from Table 4 using Tobit and the results (not reported) remain unchanged. The signs and levels of statistical significance for all variables are the same for the OLS and Tobit specifications.

Third, because of potential dependence among observations from the same physicians over multiple years, we reran Model 2 from Table 4 separately for each of the three survey samples, and the results (not reported) remain unchanged. We next conducted an additional check for the potential effect of multiple observations from some physician practices (we lack the data to exclude all such cases). Specifically, we reran Model 2 from Table 4 after dropping all physicians working in practices with more than two physicians. The results (not reported) continue to hold for these subsamples in which the potential effect of multiple observations from the same practices is minimized. In addition, we combined the two preceding procedures by running OLS Model 2 from Table 4 separately for each of the three survey administrations while also including only responses from physician practices with one or two physicians. Again, the results from OLS Model 2 in Table 4 continue to hold.

Fourth, we reran Model 2 from Table 4 after replacing the *MGPCI* measure of malpractice liability with *DIRLIAB* and *INDLIAB*, the measures of legal liability based on the presence or absence in each state of specific legal limitations on malpractice liability. Table 5 includes two sets of specialist-high legal liability categorical variables, one set using *HIGH-DIRLIAB* (above median *DIRLIAB*) and the other using *HIGH-INDLIAB* (above median *INDLIAB*), for each of the three types of physician specialists. However, the Specialist dummy variables (e.g., *MEDSPEC*) are highly correlated with the corresponding interaction terms (e.g., *MEDSPEC**HIGH-DIRLIAB*, *MEDSPEC**HIGH-INDLIAB*) with correlations ranging from 0.56 to 0.76. As a result, the interaction terms in Model 2 Table 5 have estimated coefficients that are not statistically different from zero. Results for Model 2A in the last column of Table 5 show that rerunning Model 2 after dropping the Specialist dummy variables yields results that are qualitatively similar to, but weaker than, the results for Model 2 of Table 4.

Finally, we checked two further specifications. First, we introduced additional controls for whether the individual physician received bonuses based on quality, patient satisfaction, or cost control measures, as well as the total percentage of the physician's income received in the form of bonuses. Second, we constructed an alternative measure of the extent of risk imposed on physicians by summing the practice's percentage of capitation revenue (*PCAPREV*) and the individual physician's percentage of bonus compensation. In both

TABLE 5
Effect of Physician Specialty and Malpractice Costs (*HIGH-DIRLIAB* and *HIGH-INDLIAB*) on Physician Practice's Capitation Revenue for Surveys 1–3[a] (OLS Estimates)

Dependent variable: log (1 + *PCAPREV*)

Independent Variables	Model 2 with Direct Liability/ Indirect Liability Coefficient (t^b)		Model 2A with Direct Liability/ Indirect Liability Coefficient (t^b)	
Intercept	−3.17	−23.32**	−1.84	−16.13**
FAM/GEN	1.32	16.08**		
INTERNAL	1.32	15.89**		
PED	1.21	14.26**		
MEDSPEC	0.37	4.22**		
SUGSPEC	0.16	1.82		
PMC	0.02	44.78**	0.02	48.37**
log (*SIZE*)	0.18	21.45**	0.18	20.83**
EFGUID	0.13	16.95**	0.13	16.40**
log (*POPULATN*)	0.19	23.99**	0.17	20.88**
S2 DUMMY	−0.05	−1.98	−0.08	−3.36**
S3 DUMMY	−0.34	−13.87**	−0.34	−13.65**
*MEDSPEC*HIGH-DIRLIAB*	−0.07	−1.30	−0.17	−3.05**
*SUGSPEC*HIGH-DIRLIAB*	0.02	0.45	−0.11	−2.13
*OB/GYN*HIGH-DIRLIAB*	0.13	1.08	−0.11	−0.94
*MEDSPEC*HIGH-INDLIAB*	0.17	3.11	−0.46	−10.13**
*SUGSPEC*HIGH-INDLIAB*	−0.03	−0.53	−0.86	−20.28**
*OB/GYN*HIGH-INDLIAB*	−0.09	−0.74	−1.04	−11.20**
Adjusted R^2	.30		.25	
n	18,680		18,680	

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*, ** Indicate two-tailed significance at the 0.01 and 0.001 levels, respectively.

[a] The 1996–1997 survey, the 1998–1999 survey, and the 2000–2001 survey collected data as of 1995, 1997, and 1999, respectively.

[b] We report asymptotic t-statistics based on White (1980) standard errors.

PCAPREV = percent of practice's patient care revenue paid on capitated or other prepaid basis;
FAM/GEN = family/general practice;
INTERNAL = internal medicine;
PED = pediatrics;
MEDSPEC = medical specialties;
SUGSPEC = surgical specialties;
OB/GYN = obstetrics and gynecology;
DIRLIAB = malpractice liability as reduced by reform laws that *directly* limit the amount of damage award to patient(s), larger values indicate greater legal liability;
HIGH-DIRLIAB = a dummy variable for high *DIRLIAB* coded as 1 for the above median *DIRLIAB*, and 0 for the below median *DIRLIAB*;
INDLIAB = malpractice liability as reduced by reform laws that *indirectly* limit the frequency of malpractice litigation, larger values indicate greater legal liability;
HIGH-INDLIAB = a dummy variable for high *INDLIAB*, coded the same way as *HIGH-DIRLIAB*;
PMC = percentage of practice's patient care revenue from all managed care organizations (MCOs);
SIZE = the number of physicians working for the practice;
EFGUID = extent to which formal/written practice guidelines from HMO and medical group affect the physicians' practice (0 = no effect, 1 = very small, 2 = small, 3 = moderate, 4 = large, and 5 = very large);
POPULATN = population of the Metropolitan Statistical Area (MSA) in which physician practice is located;
S2 DUMMY = an indicator variable for Survey 2 (1998–1999);
S3 DUMMY = an indictor variable for Survey 3 (2000–2001); and
log = natural log.

cases we use individual physician's compensation characteristics to proxy for the practice characteristics. The results (not reported) remain robust to each of these alternative specifications.

VI. DISCUSSION AND CONCLUSION

This study examines the determinants of the form of incentive contracts between MCOs and physician practices. We focus on the role of physicians' task characteristics and states' physician malpractice provisions as factors influencing the benefits and costs of imposing risk on physician practices by paying them on a capitation basis. Our results are generally consistent with the agency moral hazard analysis of the effect of uncertainty on the strength of incentive provisions in agents' contracts. Proposition 1 establishes that as cost control measures become noisier, it is optimal for MCOs to reduce the weight assigned to cost control measures relative to the weight assigned to revenue generation measures in physician contracts. Consistent with this prediction, our empirical results show that when MCOs move from a PCP's environment, in which exogenous cost risk is smaller, to a specialist's environment with greater exogenous cost risk, they significantly reduce the extent of cost risk (capitation) in physicians' contracts.

Next, we consider the effect of cross-sectional variation in the association between cost control and legal liability outcomes, as proxied by the specialist-high MGPCI dummy variables. Our model predicts that when malpractice liability is only weakly associated with cost control outcomes, increasing the level of malpractice liability damages alone should have no effect on the extent of capitation in physicians' contracts. However, when unfavorable cost control outcomes are more strongly associated with greater malpractice liability outcomes, the model predicts that specialists will be compensated less on a capitation basis. Our empirical results support this prediction.

Our conclusions regarding the central role of risk-aversion and its differential role for PCPs versus specialists is supported in the professional literature on the design of physician incentive contracts. In *The Managed Care Handbook*, Kongstvedt (2001, 126) emphasizes the importance of uncertainty and physicians' risk aversion in the design of incentive contracts for physicians:

> The most common problem with capitation involves chance ... Physicians with fewer than 100 members may find that the dice simply roll against them, and they will have members who need bypass surgery, have cancer, have acquired immune deficiency syndrome (AIDS), or a host of other expensive medical problems.

Similarly, John Steenwyk, President of Health Economics, describes why "caps" (i.e., capitation) is more effective for PCPs than for specialists:

> Except for primary care services, which involve little risk and are fairly steady, individual doctor's simply are not equipped to accept insurance risk. (That's why we license insurers.) Primary caps aren't so fatal—the real mischief is in specialty service caps. (as quoted in Stenson [2000, 63])

Our results documenting a negative relation between the extent of uncertainty and the strength of cost-sharing provisions in agents' contracts provide an interesting contrast to previous findings in other contexts. Prendergast (2002) summarizes previous evidence of a *positive* relation between the level of noise in performance measures and the extent of risk borne by the agents. Prendergast attributes these findings to the principal's desire to provide the agent with the necessary incentives to use superior local information in decisions. This private information perspective suggests that it would be reasonable to expect the same relation in our setting because physicians obviously possess extensive and important superior information about patients' conditions and the expected efficacy of alternative treatments. However, we find the opposite, i.e., a *negative* relation between the level of noise in performance measures and the extent of risk borne by the agents, as predicted by agency moral hazard models.

We attribute the inconsistency between our results on the strength of contractual incentives and the predictions from private information analyses to an important distinction between the MCO-physician contractual setting and the models developed by Prendergast (2002) and Lafontaine and Bhattacharya (1995). In their analyses, the agent has superior information about local operating conditions and the principal must decide how much discretion to grant the agent. The principal's choice is between granting an agent more discretion and then using high-powered incentives to motivate the agent to use the private information as the principal wishes, or granting less discretion and controlling the agent by directly monitoring the agent's action choices.

In our view, the latter approach of relying on input monitoring constitutes a practical alternative to incentive contracts in some environments, but typically not in the environments in which physicians practice. For example, input monitoring is more likely to be an effective alternative in the environments reviewed by Prendergast (2002), such as a franchisee operating a local retail outlet or a sharecropper working a landowner's fields. However, extensive input monitoring combined with minimal discretion seems fundamentally inconsistent with the role of an agent in situations involving extensive and complex technical insight and related discretion, such as a physician. Even PCPs, who typically face less uncertainty than do specialists, nevertheless exercise a great deal of discretion in treating patients. As a result, Prendergast's (2002) model, in which the MCO would specify in advance detailed treatment protocols for all patients, is likely to be less appropriate for our setting.

Our results are subject to several qualifications. First, our measurement of the extent of capitation is based on reports from individual physicians rather than from the actual records of physician practices. To the extent that physicians' reports are noisy measures, this should operate against our finding significant results. Likewise, we measure the extent of capitation in contracts between MCOs and physician practices, whereas we noted earlier that some practices include a second layer of contracts between the physician practices and individual physicians. We assume that such practices operating under capitation contracts will transmit corresponding incentives to individual physicians. However, to the extent that they do not, this could again operate against our finding empirical support for our hypotheses.

Second, our tests of the effect of malpractice liability pressure would ideally employ time-series data on the evolution of MCO-physician practice contracts as a function of changes in the legal liability environment in each state. Such time-series data would be more effective in controlling for the effect of correlated omitted variables and for the influence of factors such as state legal provisions that may be endogenous. However, our time-series data include only three observations over six years, and hence we must rely primarily on cross-sectional tests.

Finally, we treat the extent of delegation of decisions by MCOs to physicians as exogenous, but more generally principals will choose the extent of delegation and the strength of incentives simultaneously as in Nagar (2001). This issue may be less important in our setting to the extent that all physicians exercise a relatively high level of discretion, in which case our sample would reflect less variation in the level of delegation. Consistent with this interpretation, the results in Table 1 for *EFGUID*, the relative effect of practice guidelines across physician specialties, show that the guidelines have limited effects, and these effects remain relatively constant across specialties. Finally, another constraint on our ability to explain the level of delegation for our sample is that, compared to Nagar's (2001) four broad measures of delegation, our single measure is likely to be significantly noisier.

APPENDIX A

Specialist Capitation

MCOs initially employed capitation in contracts with PCPs, and later extended capitation to contracts with specialists. An example of specialist capitation developed by Horowitz (1997) illustrates not only two specialist capitation methods, but also provides insight to the costs and benefits of specialist capitation.

Horowitz (1977) begins with a *traditional capitation contract* for a neurosurgeon under which the MCO pays the neurosurgeon \$0.10 PMPM, meaning \$0.10 per each covered member per month. Actuaries calculate this capitation rate based on experience and judgment, estimating that for every 1,000 enrollees, two will be referred to a neurosurgeon annually for a referral rate of 0.002. Similarly, the actuaries estimate that of those referred, 80 percent will be treated without surgery at a cost to the neurosurgeon of \$250 per enrollee. The other 20 percent will require surgery at a cost to the neurosurgeon of \$2,000 per enrollee. Thus, the neurosurgeon's expected annual cost per 1,000 enrollees is $[.002 \times 1{,}000] \times [.8 \times \$250 + .2 \times \$2{,}000] = \$1{,}200$. The capitation rate of \$0.10 PMPM then pays the neurosurgeon $[1{,}000] \times \$0.10 \times 12$ months $= \$1{,}200$, which equals the neurosurgeon's expected cost of patient care. The neurosurgeon would receive a predictable cash flow of \$100 per month starting when the contract became effective. On the other hand, the neurosurgeon would bear the risk of the number of referred patients as well as the cost per referred patient.

Alternatively, some MCOs employ *contact capitation* contracts in which payments do not begin until a patient is referred to the neurosurgeon, i.e., until the first "contact" with the patient. For the above example, the contact capitation rate would be \$1,200 of expected cost divided by two expected patients, yielding \$600 per contact. Thus, the neurosurgeon would receive \$600 each time treatment began on a new patient. The neurosurgeon would bear the risk of whether the cost of treating the patient would be \$250, \$2,000, or some other amount.

Some specialty capitation arrangements employ a risk-sharing procedure whereby, for each physician specialty, all monthly capitation payments received by all physician groups practicing under the plan are placed into a single "risk pool," or alternatively multiple risk pools, one for commercial insurance, one for Medicare, and one for Medicaid. The risk-pool covers the costs of a contractually defined set of services, typically including inpatient and outpatient professional services, office surgeries, interpretation of radiology services, office diagnostic testing, injectables, and office visits (Frank and Brunsberg 1999). Examples of costs frequently not included are implantable devices, facility overhead costs, out-of-network costs, and prescription drugs.

Individual physicians then receive payments from the pool that in some cases are equal for all physicians and in other cases are proportional to the number of patients treated (the contact capitation method) or to particular procedures performed. Dividing the pool equally among all specialists offers risk-sharing advantages,[21] but it fails to reward physicians who are more productive by attracting more patients or performing more procedures. Contact capitation has desirable incentive features, but creates significant risk for the individual physician because of the uncertain number of referrals in any period. Basing payments from the pool to individual physicians on the number of procedures performed is similar to fee-for-service (FFS) arrangements and creates the same incentive to provide excess services as FFS does.

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[21] Gaynor et al. (2004) discuss the costs and benefits of such procedures, including the benefit of allowing a physician with particularly needy patients in one year to share the cost of treatment with other physicians.

Various other contractual arrangements are employed with risk pools, including payments of bonuses or "withholds" based on the level of utilization, payment of quality bonuses, etc. Further, additional adjustments can be made to recognize variations within specialties, e.g., leukemia cases may be weighted three to four times higher than other oncology cases (Frank and Brunsberg 1999), or cost "subpools" can be created to treat various subcategories of care individually. However, such refinements significantly increase the administrative costs of such systems.

An important administrative consideration for MCOs in capitating specialists is that a specialist requires a larger panel (set of patients) than a PCP does. While a PCP may operate reasonably with a panel of 400 to 600 patients, a specialist requires approximately triple that number in order to avoid problems of excessive risk because of too small a sample size (Kongstvedt 2001, 181). This issue is related to the relatively greater reliance on capitation for PCPs than for specialists, but the source of the effect involves a combination of risk-aversion of the physician and administrative costs.

APPENDIX B
Proofs

Proof of Proposition 1

From Equation (5), the first-order conditions for the MCO's choice of contract parameters β_1 and β_2 are:

$$f_1 g_{11} - [\beta_1 g_{11} - L_1] g_{11} - L_1 g_{11} - r\beta_1 \text{ var}_1 = 0, \text{ and} \qquad (6)$$

$$f_2 g_{22} - [\beta_2 g_{22} - L_2] g_{22} - L_2 g_{22} - r\beta_2 \text{ var}_2 + r \text{ cov}(e_2,e_L) = 0. \qquad (7)$$

Simplifying and rearranging the first-order conditions gives:

$$\beta_1[g_{11}^2 + r \text{ var}_1] = f_1 g_{11}; \qquad (6)$$

$$\beta_2[g_{22}^2 + r \text{ var}_2] = f_2 g_{22} + r \text{ cov}(e_2,e_L). \qquad (7)$$

Solving Equations (6) and (7) simultaneously for β_1 and β_2, and rearranging yields:

$$\frac{\beta_2}{\beta_1} = \frac{\{[f_2 g_{22}][g_{11}^2] + r \text{ var}_1[f_2 g_{22}] + [g_{11}^2] r \text{ cov}(e_2,e_L) + r^2 \text{ var}_1 \text{ cov}(e_2,e_L)\}}{\{[f_1 g_{11}][g_{22}^2] + r \text{ var}_2[f_1 g_{11}]\}}. \qquad (8)$$

From Equation (6), with $r > 0$, $\text{var}_1 > 0$, $f_1 > 0$ and $g_{11} > 0$ by assumption, it must be that $\beta_1 > 0$. Likewise, from Equation (4) with $a_2 > 0$, $L_2 > 0$ and $g_{22} > 0$ by assumption, it must also be the case that $\beta_2 > 0$, consistent with the physician's payoff increasing in the cost control measure. Then with $\beta_1 > 0$ and $\beta_2 > 0$, it follows that $\beta_2/\beta_1 > 0$. Next, take the partial derivative of Equation (8) with respect to var_2 to evaluate how the relative weights on the cost and revenue measures vary with the precision of the cost measure:[22]

$$\frac{\partial(\beta_2/\beta_1)}{\partial \text{var}_2} = \\ -\frac{\{[f_2 g_{22}][g_{11}^2] + r\,\text{var}_1[f_2 g_{22}] + [g_{11}^2]r\,\text{cov}(e_2,e_L) + r^2\,\text{var}_1\,\text{cov}(e_2,e_L)\}\{r[f_1 g_{11}]\}}{\{[f_1 g_{11}][g_{22}^2] + r\,\text{var}_2[f_1 g_{11}]\}^2}. \tag{9}$$

From Equation (8), the sum of the four terms in the first braces in the numerator of Equation (9) is strictly positive because the sum equals β_2, which was shown above to be strictly positive. Likewise, by assumption we have $r > 0$, $f_1 > 0$, $g_{11} > 0$, which implies that $\{r[f_1 g_{11}]\} > 0$. Hence, with both braced terms in the numerator of Equation (9) strictly positive, the numerator is positive. The denominator is positive by inspection. Together with the sign on the fraction in Equation (9), the positive numerator and denominator imply that, $\frac{\partial(\beta_2/\beta_1)}{\partial \text{var}_2} < 0$, which establishes Proposition 1.

Proof of Proposition 2

Rewrite Equation (8) as:

$$\frac{\beta_2}{\beta_1} = \frac{\{[f_2 g_{22}][g_{11}^2] + r\,\text{var}_1[f_2 g_{22}]\} + \{g_{11}^2\,r + r^2\,\text{var}_1\}\text{cov}(e_2,e_L)}{[f_1 g_{11}][g_{22}^2 + r\,\text{var}_2]} \tag{10}$$

which can then be written as:

$\beta_2/\beta_1 = Q_0 + Q_1\,\text{cov}(e_2,e_L)$, where Q_0 and Q_1 are positive constants.

Substituting $\text{cov}(e_2,e_L) = \rho_{2L}\,\text{var}_2\,\text{var}_L$, gives $\beta_2/\beta_1 = Q_0 + Q_1(\rho_{2L}\,\text{var}_2\,\text{var}_L)$, which can be written as $\beta_2/\beta_1 = Q_0 + Q_2\rho_{2L}$, with $Q_0, Q_2 > 0$.

Taking the partial of β_2/β_1 for the optimal ratio of incentive weights with respect to ρ_{2L} yields $\frac{\partial(\beta_2/\beta_1)}{\partial \rho_{2L}} > 0$, which is the desired result. Given $\rho_{2L} < 0$ by assumption, an *increase* in ρ_{2L} would make it closer to zero and such a weaker association would result in more capitation. Then conversely, again because $\rho_{2L} < 0$, a *decrease* in ρ_{2L} would mean the negative association between cost control and malpractice liability would become stronger, and Proposition 2 would predict that the extent of capitation would *decrease*.

REFERENCES

Banker, R., and S. Datar. 1989. Sensitivity, Precision and Linear Aggregation of Signals for Performance Evaluation. *Journal of Accounting Research* 27 (1): 21–39.

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[22] We note that if $\text{cov}(e_2,e_L) = 0$, Equation (8) reduces to Equation (3) from Datar et al. (2001), and hence their interpretation of the optimal weights on the incentive measures y_1 and y_2 in terms of the sensitivity and precision of each measure would apply here. However, because we expect a negative association between the cost-control effectiveness (y_2) and legal liability exposure (L), i.e., $\text{cov}(e_2,e_L) < 0$, our analysis in Proposition 2 focuses on the effect of changes in the level of this association.

Bazzoli, G., L. Dynan, and L. Burns. 1999. Capitated contracting of integrated health provider organizations. *Inquiry* 36 (Winter): 426–444.

Brown, V., and B. Hartung. 1998. Managed care at the crossroads: Can managed care organizations survive government regulation? In *Annuals of Health Law, Loyola University Chicago Institute for Health Law*, 7, Ann. Health L. 25. Chicago, IL: The Institute for Health Law, Loyola University.

Butler, P. 2001. *Key Characteristics of State Managed Care Organization Liability Laws: Current Status and Experience*. Menlo Park, CA: The Henry J. Kaiser Family Foundation.

Center for Studying Health System Change. 2002. *Community Tracking Study Physician Surveys 1996–97, 1998–99 and 2000–01*. United States, computer file, ICPSR version. Produced 2001: Washington, D.C.: Center for Studying Health System Change. Distributed 2002: Ann Arbor, MI: Inter-University Consortium for Political and Social Research.

Datar, S., S. Kulp, and R. Lambert. 2001. Balancing performance measures. *Journal of Accounting Research* 39 (1): 75–92.

Davenport, J. 2000. Documenting high-risk cases to avoid malpractice liability. *Family Practice Management* 7 (9).

Evans, J. H., A. Leone, and N. J. Nagarajan. 2005. Non-financial performance measures in the healthcare industry: Do quality-based incentives matter? *Advances in Management Accounting* 14.

Feltham, G. A., and J. Xie. 1994. Performance measure congruity and diversity in multi-task principal/agent relations. *The Accounting Review* 69 (3): 429–453.

Frank, C., and J. Brunsberg. 1999. Using contact capitation to align payment incentives among specialists. *Healthcare Financial Management* 53 (10): 52–56.

Gaynor, M., J. Rebitzer, and L. Taylor. 2004. Physicians incentives in HMOs. Working paper, NBER and Carnegie Mellon University.

Gal-Or, E. 1999. Optimal reimbursement and malpractice sharing rules in health care market. *Journal of Regulatory Economics* 16: 237–265.

Gold, M., L. Nelson, T. Lake, R. Hurley, and R. Berenson. 1995. Behind the curve: A Critical assessment of how little is known about arrangements between managed care plans and physicians. *Medical Care Research and Review* 52 (3): 307–341.

Harvard Medical Practice Study, Patient, Doctors, Lawyers. 1990. *Medical Injury, Mal-Practice Litigation and Patient Compensation in New York*. Cambridge, MA: Harvard University.

Health Care Financing Administration. 2001. *42 CFR Parts 410 and 414, Medicare Program: Revisions to Payment Policies under the Physician Fee Schedules for Calendar Year 2001*. HCFA-1120-P. Washington, D.C.: Office of the Federal Register.

Hillman, A. 1987. Financial incentives for physicians in HMOs: Is there a conflict of interest? *New England Journal of Medicine* 317.

———, W. Welch, and M. Pauly. 1992. Contractual arrangements between HMOs and primary care physicians: Three-tiered HMOs and risk pools. *Medical Care* 30 (2): 136–148.

Holmstrom, B. 1979. Moral hazard and observability. *Bell Journal of Economics* 10: 74–91.

Horowitz, J. 1997. Contact capitation: An alternative for specialists capitation. *Healthcare Financial Management* 51 (11): 54–55.

Ittner, C., D. Larcker, and M. Pizzini. 2006. Performance-based compensation in professional service firms: An examination of medical group practices. Working paper, University of Pennsylvania, Stanford University, and Southern Methodist University.

Katon, W., M. Von Korff, E. Lin, and G. Simon. 2001. Rethinking practitioner roles in chronic illness: The specialists, primary care physicians, and the practice nurse. *General Hospital Psychiatry* 23: 138–144.

Kessler, D., and M. McClellan. 1996. Do doctors practice defensive medicine? *Quarterly Journal of Economics* (May): 353–390.

———, and ———. 2000. Medical liability, managed care, and defensive medicine. Working Paper No. 191, Stanford Law School, John M. Olin Program in Law and Economics.

Kongstvedt, P. 2001. *The Managed Health Care Handbook*. 4th edition. Gaithersburg, MD: Aspen Publishers, Inc.

Lafontaine, F., and S. Bhattacharya. 1995. The role of risk in franchising. *Journal of Corporate Finance* 2: 39–74.

Leone, A. 2002. The relation between efficient risk sharing arrangements and firm characteristics: Evidence from the managed care industry. *Journal of Managerial Accounting Research* 14: 99–117.

McCullough, Campbell & Lane LLP. 2003. *Summary of Medical Malpractice Laws*. Available at: http://www.mcandl.com/states.html.

Meyer, G., E. Cheng, and J. Elting. 2000. Differences between generalists and specialists in characteristics of patients receiving gastrointestinal Procedures. *Journal of General Internal Medicine* 15 (3): 188–194.

Nagar, V. 2002. Delegation and incentive compensation. *The Accounting Review* 77 (2): 379–395.

Phelps, E. C. 1997. *Health Economics*. 2nd edition. Reading, MA: Addison-Wesley Educational Publishers Inc.

Prendergast, C. 2002. The tenuous trade-off between risk and incentives. *Journal of Political Economy* 110: 1071–1102.

Stenson, R. 2000. *A Physician's Guide to Thriving in the New Managed Care Environment: Selecting the Right Strategy for Your Practice*. Wilsonville, OR: BookPartners, Inc.

U.S. Census Bureau. 2000. *Census*. Redistricting Data (P.L. 94-171) Summary File and 1990 Census. Available at: www.census.gov/population/cen2000/phc-t3/tab02.pdf.

U.S. Congress, Office of the Technology Assessment. 1993. *Impact of Legal Reforms on Medical Malpractice Costs*. OTA-BP-H-119. Washington, D.C.: Government Printing Office.

White, H. 1980. A heteroscedasticity-consistent covariance matrix estimator and a direct test for heteroscedasticity. *Econometricia* 48 (4): 817–838.

Zuckerman, S., and S. Norton. 1994. *Updating the Geographic Practice Cost Index: The Malpractice GPCI*. The Urban Institute, Health Care Policy Center Sponsored by Health Care Financing Administration, PB94-161106. Baltimore, MD: Health Policy Center.

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The Evolution of Stock Option Accounting: Disclosure, Voluntary Recognition, Mandated Recognition, and Management Disavowals

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ABSTRACT: In this study we report the results of an experiment that examines how relatively sophisticated financial statement users interpret management stock option compensation disclosures under SFAS No. 123 and SFAS No. 123R. We predict and find that mandated income statement recognition, as required under SFAS No. 123R, leads to higher user assessments of reliability than either voluntary income statement recognition or voluntary footnote disclosure, options allowed under SFAS No. 123. Users view voluntary footnote disclosure as the least reliable reporting alternative. We also examine the amount users invest in response to these accounting treatments, and find that users invest more in a firm when management chooses income statement recognition than when management chooses footnote disclosure. We find no difference in investment amounts between mandated recognition and either voluntary recognition or footnote disclosure. Finally, although we find that these results are insensitive to whether management explicitly disavows the reliability of stock option expense, we present evidence that, in side-by-side comparisons, where one firm disavows and the other does not, disavowals may affect user judgments and decisions.

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I. INTRODUCTION

Accounting for employee stock options is a contentious and controversial topic (Botosan and Plumlee 2001; Revsine et al. 2002; Hughes 2005). The Financial Accounting Standards Board (FASB) advocates income statement recognition (FASB 1995), whereas opponents of expensing stock options claim that the methods used to estimate stock option expense are unreliable (Malkiel and Baumol 2002). When the FASB issued Statement of Financial Accounting Standard (SFAS) No. 123, *Accounting for Stock-Based Compensation,* in 1995, almost every firm used the discretion allowed in SFAS No. 123 to disclose stock option expense in the footnotes rather than recognize it on the income statement, even though the FASB strongly recommended income statement recognition. In many cases, management has even included language in the footnotes explicitly disavowing the methods required by the FASB to measure stock option expense. Blacconiere et al. (2004) report that approximately 18 percent of the footnote-disclosing firms in their sample disavow the stock option disclosure, typically stating that option-pricing models generate an unreliable number for stock option expense.[1]

In the wake of public concerns about the transparency and quality of financial reporting that followed the Enron and WorldCom scandals, some firms voluntarily switched from footnote disclosure to income statement recognition. The FASB also revisited the issue of stock option accounting, issuing SFAS No. 123R, *Share Based Payment,* in December 2004. This new standard replaces SFAS No. 123 and mandates income statement recognition. Lobbying against required recognition of stock option expense was once again intense (Crooch 2005; Hughes 2005).

In this study we report the results of an experiment that examines how four stock option accounting issues affect the judgments and decisions of relatively sophisticated financial statement users (hereafter users). The first issue is whether users perceive differences in the reliability of stock option expense voluntarily recognized by management (as allowed under SFAS No. 123) versus recognized due to a FASB mandate (as required under SFAS No. 123R).[2] This question introduces a new issue to the accounting literature—whether the source of the decision to recognize an item on the face of the financial statements (management's choice versus a FASB mandate) influences user reliability assessments. It also provides evidence on the consequences associated with the decision to mandate or allow management discretion in applying accounting principles, a fundamental issue facing standard setters, and as Barton and Waymire (2004, 69) state, "an important issue for future research." We predict and find that users assess stock option expense recognized due to a FASB mandate to be more reliable than stock option expense voluntarily recognized by management.

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[1] For example, in its 2003 annual report, Intel states in its "Employee Stock Benefit Plans" footnote that, "SFAS No. 123 requires the use of option pricing models that were not developed for use in valuing employee stock options. The Black-Scholes option pricing model was developed for use in estimating the fair value of short-lived exchange-traded options ... Because the company's employee stock options have characteristics significantly different from those of traded options, and because changes in subjective input assumptions can materially affect the fair value estimate, in the opinion of management, the existing models do not necessarily provide a reliable single measure of the fair value of employee stock options."

[2] Reliability is an important component of the FASB's Conceptual Framework (FASB 1980, 1984), and many of the arguments against expensing stock options center on the reliability of the expense number generated by option-pricing models.

The second issue focuses on the discretion allowed under SFAS No. 123, and examines whether users perceive differences in the reliability of stock option expense when a firm voluntarily recognizes the expense versus discloses it. This question examines an issue that not only is fundamentally important to managers, accountants, and standard setters (Barth et al. 2003; Davis-Friday et al. 2004; Libby et al. 2005), but has also puzzled academic accountants (e.g., see Bernard and Schipper 1994; Barth et al. 2003; Cotter and Zimmer 2003; Davis-Friday et al. 2004): *Why* do users treat recognized and disclosed items differently? Although prior research provides evidence consistent with users treating recognized and disclosed items differently (e.g., see Harper et al. 1987; Imhoff et al. 1993; Davis-Friday et al. 1999; Espahbodi et al. 2002; Hirst et al. 2004; Hodge et al. 2004), and suggests that disclosed items are less value-relevant than recognized items (Davis-Friday et al. 1999; Cotter and Zimmer 2003), little is known about why recognition versus disclosure affects users. Archival studies face design and econometric issues that make it difficult to clearly demonstrate why (Bernard and Schipper 1994). We predict and find that users consider stock option expense voluntarily recognized on the income statement to be more reliable than stock option expense disclosed in the footnotes, suggesting that differences in perceived reliability might partially explain why users treat recognized and disclosed items differently.

The third issue we examine is whether user investment decisions differ across mandated income statement recognition, voluntary income statement recognition, and footnote disclosure of stock option expense. Managers, accountants, and standard setters are ultimately interested in how accounting treatments and disclosures affect investment behavior, and there is evidence that users may interpret recognition versus disclosure not only as a signal about the reliability of an item, but also as a signal about the firm's future prospects (Bagnoli and Watts 2005; Elayan et al. 2005; Levine and Hughes 2005). After controlling for user reliability assessments, we predict and find that users invest more in a firm that voluntarily recognizes stock option expense than in a firm that discloses the expense in the footnotes, even though voluntary recognition reduces reported net income. This finding suggests that relatively sophisticated users do not fixate on reported net income, but instead view management's choice to report conservatively as a signal of strong expected performance (Wahlen 1994; Aboody et al. 2004; Bagnoli and Watts 2005; Levine and Hughes 2005; Elayan et al. 2005). We find no differences in investment amounts between mandated recognition and either voluntary recognition or footnote disclosure of stock option expense.

The final issue we examine is whether management's choice to disavow stock option expense affects user reliability assessments and investment decisions. We are aware of no evidence concerning the effects of disavowals, and in particular stock option disavowals, on user judgments and/or decisions. Disavowals are a unique form of voluntary disclosure because they are located in audited financial statements, are provided in addition to the mandated information, and explicitly question the reliability or usefulness of the mandated information.[3] In our main analysis, we find that the presence/absence of a disavowal does not influence user reliability judgments or investment decisions in mandated income statement recognition, voluntary income statement recognition, or footnote disclosure environments. In supplemental tests where management's disavowal is more transparent, however, we provide some evidence that disavowals reduce user assessments of the reliability of stock option expense, their perceptions of the firm's "true" total compensation expense, and the amount they would invest in the firm.

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[3] Firms also disavow other mandated disclosures. For example, in its 2003 Form 10-K, Caterpillar Inc. disavows its segment footnote disclosure stating, "we feel that the financial information required by SFAS No. 131, *Disclosures about Segments of Enterprise and Related Information*, has limited value for our external readers."

Overall, the results of this study are important because they provide information about the consequences of a controversial FASB standard that eliminates management discretion and mandates placement. Our results demonstrate that both the party determining placement (management or the FASB) and placement itself (footnote or financial statements) influence user reliability assessments. The latter provides evidence consistent with the conjecture that one reason users treat recognized and disclosed items differently is that recognition signals that the recognized item meets or exceeds a threshold level of reliability (e.g., see, Bernard and Schipper 1994). In addition, we show that discretionary conservative accounting (voluntary stock option expense recognition) results in users investing more in the reporting firm, even though voluntary recognition reduces reported earnings. Finally, although our results provide only limited evidence that management disavowals of stock option expense affect user judgments, we suggest that policy makers monitor the effects of management disavowals because effective disavowals could undermine the establishment of reporting standards that provide users with comparable and consistent financial information.[4] In addition, disavowals may become even more common given that SFAS No. 123R mandates fair-value recognition of stock option expense, which management has forcefully opposed.

The remainder of the paper is organized into five sections. We present background information in Section II, discuss theory in Section III, describe the experiment in Section IV, and present and discuss the results in Sections V and VI, respectively.

II. BACKGROUND

Stock Option Accounting

Stock option accounting has been, and continues to be, an important and controversial issue (Hughes 2005). Prior to the issuance of SFAS No. 123 in 1995, Accounting Principles Board (APB) Opinion No. 25, *Accounting for Stock Issued to Employees*, governed the accounting treatment for employee stock options (APB 1972). This opinion required companies to use the intrinsic method. Under this method, a firm recognizes as an expense the excess on the grant date of the market value of its stock over the option exercise price; this amount typically is zero.

The FASB issued the exposure draft of SFAS No. 123 in 1993. The FASB proposed income statement recognition for stock option compensation based on the fair value of the options on their grant date, estimated using an option-pricing model. However, in the face of intense lobbying and the threat of Congressional intervention, the FASB relaxed its position. SFAS No. 123 still required companies to use the fair value of the options on their grant date to calculate employee stock option expense; however, the standard gave companies a choice between income statement recognition and footnote disclosure. With the former, a company recognized stock option expense as a determinant of reported earnings on its income statement. With the latter, the company reported no expense on its income statement (consistent with the intrinsic method under APB Opinion No. 25), but disclosed in its footnotes pro forma earnings that included stock option expense as a component. Although SFAS No. 123 encouraged recognition, the vast majority of U.S. firms chose footnote disclosure. Following the collapse of Enron and WorldCom, a substantial minority of firms changed their accounting treatment and started voluntarily recognizing stock option expense (Aboody et al. 2004).

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[4] In Statement of Financial Accounting Concepts (SFAC) No. 2, *Qualitative Characteristics of Accounting Information*, the FASB states that comparability is achieved if similar transactions are accounted for similarly and different transactions are accounted for differently (FASB 1980). In the same statement, the FASB notes that creating comparable information is a fundamental rationale for standard setting.

Accounting standard setters recently revisited the issue of employee stock options. The International Accounting Standards Board issued International Financial Reporting Standard No. 2, which requires firms to recognize the cost of providing stock options as an expense on their income statements beginning in January 2005. In December 2004, the FASB issued SFAS No. 123R, which requires income statement recognition of employee stock options effective June 2005. The lobbying against required recognition of stock option expense was again intense (Crooch 2005), and the Securities and Exchange Commission (SEC) postponed the implementation date of SFAS No. 123R by up to six months (SEC 2005).[5]

Research on Recognition, Disclosure, and Reliability

Income statement recognition versus footnote disclosure is a fundamental issue in accounting (Bernard and Schipper 1994; Barth et al. 2003; Davis-Friday et al. 2004; Libby et al. 2005). In general, prior research provides evidence that users treat recognized items differently than disclosed items (e.g., Harper et al. 1987; Imhoff et al. 1993; Davis-Friday et al. 1999; Espahbodi et al. 2002; Hirst et al. 2004), and that users find recognized items to be more value-relevant than disclosed items (e.g., Davis-Friday et al. 1999; Cotter and Zimmer 2003).

The literature suggests three potential reasons for why users treat recognized items differently than disclosed items: (1) consistent with the FASB's Conceptual Framework, the decision to recognize versus disclose may provide a signal about the reliability of the item (Bernard and Schipper 1994); (2) users may have greater difficulty acquiring and/or processing items disclosed in footnotes (Bernard and Schipper 1994; Hodge et al. 2004; Graham et al. 2005); and (3) because contracts typically are based on recognized amounts, recognition affects a firm's contracting costs (Bernard and Schipper 1994; Espahbodi et al. 2002). We consider the first reason in this study.[6]

Two recent cross-sectional, archival studies provide evidence consistent with users considering recognized items to be more reliable than disclosed items. Using a sample of Australian firms, Cotter and Zimmer (2003) conclude that reliability differences explain why voluntarily recognized asset revaluations are more value-relevant than voluntarily disclosed asset revaluations, but their conclusion is based on only indirect evidence. Davis-Friday et al. (2004) conclude that the market perceives disclosed obligations for post-retirement benefits other than pensions to be less reliable than recognized obligations. However, their study examines footnote disclosure relative to *mandated* recognition, not to *voluntary* recognition. Thus, their study cannot distinguish whether the difference in perceived reliability is due to (1) footnote disclosure versus voluntary recognition and/or (2) voluntary recognition versus mandated recognition. A further limitation of Davis-Friday et al. (2004) is that key variables in their tests are difficult to control for and their sample is subject to self-selection bias.[7] Bernard and Schipper (1994) argue that such problems, as well as difficult econometric issues, are inherent in archival studies that examine recognition versus disclosure. They also note that few situations exist where recognition and disclosure are true alternatives. By using an experiment (coupled with our stock option context), we are able to circumvent these concerns and provide direct evidence on whether users perceive recognized items to be more reliable than disclosed items.

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[5] During a speech at the Financial Reporting Conference in May 2005, FASB board member Michael Crooch reported that the FASB has received over 14,000 comment letters related to stock option accounting (Crooch 2005). Revsine et al. (2002) note that when the FASB was debating the expensing of stock options in the early 1990s, a Senate bill was introduced that would have eliminated the FASB's independence by requiring the Securities and Exchange Commission to approve all new FASB standards.

[6] Examining all three explanations is beyond the scope of this paper. With respect to the acquisition/processing explanation, Hodge et al. (2004) demonstrate that users who use search-facilitating technology are more aware of recognition/disclosure differences than users who do not use the technology. In addition, research suggests that user sophistication affects the acquisition of information recognized in the financial statements (see, e.g., Hirst and Hopkins 1998; Maines and McDaniel 2000), and that the placement of information within the financial statements affects how users process the information (e.g., Hopkins 1996; Maines and McDaniel 2000). With respect to the contracting cost explanation, Espahbodi et al. (2002) assert that their finding of negative (positive) abnormal returns surrounding events that increased (decreased) the likelihood of firms being required to recognize stock option expense is consistent with the contracting costs explanation.

III. HYPOTHESIS DEVELOPMENT

Under SFAS No. 123 management could disclose stock option expense in the footnotes or voluntarily recognize it as an expense on the income statement. Under SFAS No. 123R management is mandated to recognize the expense on the income statement. In this section we consider whether user assessments of the reliability of stock option expense and the amount users invest in the reporting firm differ across these three reporting environments (i.e., footnote disclosure, voluntary recognition, and mandated recognition). We argue that the three environments provide different signals to users about both the characteristics of the item being accounted for and the future prospects of the firm. We also consider whether management's choice to disavow stock option expense information affects user reliability assessments and investment behavior.

Reporting Environment and Reliability Assessments

In this section we argue that users evaluate accounting treatments by considering two factors: (1) the extent to which the characteristics (e.g., the reliability) of the resulting information meet the objectives of financial reporting, and (2) the consequences to management of the treatment. The relative importance users place on each factor varies depending on the party determining the treatment (e.g., the FASB versus management) and the nature of the treatment (e.g., disclosure versus recognition). We predict that users consider information resulting from a FASB mandate to be highly reliable because the FASB's conceptual framework and objectives are expressed only in terms of the first factor. In contrast, users consider voluntary accounting treatments by management to be at least partially driven by the second factor (Fuller and Metcalf 1978; Hodge 2003), even when the treatment is inconsistent with management's *apparent* reporting incentives. In voluntary reporting environments users are never completely certain what incentives may have influenced management's accounting choices.[8]

Research has shown that financial statement users consider how a firm accounts for an item as a signal about the characteristics of that item (e.g., Hopkins 1996; Davis-Friday et al. 1999; Cotter and Zimmer 2003; Hirst et al. 2004). A FASB mandate to recognize an item (e.g., stock option expense) on the income statement, for example, signals to financial statement users that an independent, well-informed, authoritative body believes that the item meets the necessary criteria for income statement recognition. The goal of the FASB, as stated in SFAC No. 5, is to provide information "useful to those making investment and credit decisions." Reliability is a key attribute of useful information (FASB 1980, 1984).

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[7] In Davis-Friday et al. (2004), the amount and nature of supplemental information differed between the recognition and disclosure groups, and in many cases the disclosed obligation was a range estimate while the recognized obligation was a point estimate. These issues provide potential alternative explanations for their observed results.

[8] For example, management might voluntarily recognize stock option expense because it has incentives to signal strong future financial performance (Hughes and Levine 2005) or transparent financial statements.

An item must, therefore, meet a relatively high threshold level of reliability to warrant mandated recognition on the face of the financial statements. Further, the controversial nature of stock option accounting and the protests endured by the FASB provide further evidence to users that the FASB strongly believes that this expense and the methods used to calculate it are sufficiently reliable to justify income statement recognition by all publicly traded firms.

What about voluntary accounting choices made by management? Users view such choices as motivated by some combination of management's incentive to report strategically and management's desire to provide relevant and reliable (i.e., useful) information (Fuller and Metcalf 1978; Hodge 2003). Absent certainty about what motivated management's choice, users are concerned that management made the choice in part to achieve some biased reporting objective, not because it viewed the item as reliable or unreliable. We therefore expect users to view information as more reliable when its treatment is mandated by the FASB than when its treatment is voluntarily chosen by management; users view FASB mandates to be determined by the characteristics of the information and not by the consequences to management of the accounting treatment.

Users, however, do not view all voluntary accounting choices as equally strategic. When management has a choice about how to account for an item (e.g., the choice to disclose or recognize stock option expense under SFAS No. 123), users may view the choice itself as a signal about the degree of management bias. Attribution and persuasion theory (Kelley 1972; Eagly and Chaiken 1975; Fiske and Pavelchak 1986) suggests that the degree of bias users perceive in information reported by management depends on the consistency between the reported information and the users' beliefs about management's reporting incentives. If users believe that management faces incentives to be optimistic, and management reports incentive-consistent, optimistic information (confirming expectations), attribution theory suggests that users will view this information as more strategic and more biased than if management had reported incentive-inconsistent information. Users view incentive-consistent information as self-serving, intended to promote management's position. In contrast, users view incentive-inconsistent information as less biased because such information appears less self-serving.[9] Hirst et al. (1995) and Hodge et al. (2006) provide empirical evidence supporting this view, documenting that users perceive information consistent with a firm's incentives to be more biased than information inconsistent with its incentives. Archival studies examining the disclosures of financially distressed firms (Frost 1997; Koch 1999) further corroborate that management incentives influence user assessments.

Choosing to recognize stock option expense on the income statement reduces reported net income while choosing to disclose the expense in the footnotes does not. Given that managers generally have incentives to report higher net income, attribution theory suggests that users will view voluntary recognition as a less biased reporting choice than disclosure. Less perceived bias implies that users will view management's decision to recognize the expense as motivated less by strategic incentives and more by the desire to report relevant and reliable (useful) information. This line of reasoning is further supported by others who argue that recognized items have a higher reliability threshold than do disclosed items (FASB 1980; Bernard and Schipper 1994; Libby et al. 2005).

The discussion above suggests that user reliability assessments of stock option expense will vary across our three reporting environments in the following manner:

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[9] This view is consistent with signaling theory (Spence 1974; Titman and Trueman 1986; Bagnoli and Watts 2005; Levine and Hughes 2005). Under this theory, users perceive a signal to be credible only if the signal is costly to the firm. A signal that is inconsistent with the firm's incentives is costly to the firm.

H1: User reliability assessments of stock option expense will decrease across the following reporting environments: mandated income statement recognition, voluntary income statement recognition, and footnote disclosure.

Reporting Environment and Investment Decisions

In the previous section we argued that users view management's reporting of stock option expense as a signal about the characteristics of the expense. In this section we argue that users also view management's reporting of the expense as a signal about the future prospects of the firm.

Levine and Hughes (2005) and Bagnoli and Watts (2005) propose that users interpret accounting choices as signals about a firm's future prospects, and that users view firms that voluntarily choose income-reducing accounting methods as having more favorable future prospects than firms that do not. Archival research linking conservative reporting to higher stock prices (Wahlen 1994; Aboody et al. 2004; Barton and Waymire 2004; Elayan et al. 2005) provides additional support for this argument. This research suggests that users will view a voluntary choice by management to report lower income (voluntary recognition) as a relatively positive signal about the firm's future prospects and a voluntary choice by management to report higher income (voluntary disclosure) as a relatively negative signal about the firm's future prospects. Because management has no choice when a specific accounting treatment is mandated, users will view a mandated accounting environment as uninformative—and thus a neutral signal—about the firm's future prospects. To the extent that user beliefs about a firm's future prospects influence their investment decisions (Ball and Brown 1968; Beaver 1968; Nichols and Wahlen 2004), the discussion above suggests that user investment levels will vary across our three reporting environments in the following manner:

H2: The amount users invest in a firm will decrease across the following reporting environments: voluntary income statement recognition, mandated income statement recognition, and footnote disclosure.

We expect user reliability assessments of stock option expense to be negatively associated with their investment decisions. Earnings that include stock option expense are always lower than earnings that exclude the expense. Thus, as a user's perception of the reliability of stock option expense increases, that user's perception of the firm's performance, as measured by current and future earnings, should decrease. To the extent that estimates of a firm's future earnings relate positively to investment decisions (Ball and Brown 1968; Beaver 1968; Nichols and Wahlen 2004), user reliability assessments of stock option expense will be negatively associated with investment amounts.[10] Further, as predicted in H1, we expect user reliability assessments to differ across the three reporting environments. For these reasons it is important to control for user reliability assessments when testing H2.

Effect of Disavowals on Reliability Assessments and Investment Decisions

What happens to user reliability assessments and investment decisions when management includes a voluntary statement in its stock option footnote that disavows the reliability of the method used to calculate stock option expense? The effect of such a disavowal is not clear. On one hand, users may view the disavowal as an informative voluntary disclosure. Users may believe that management is using the disavowal to communicate private information that its estimate of stock option expense is unreliable, causing users to perceive stock option expense to be less reliable. In addition, research shows that the mere presence of information, regardless of its relevance to a judgment or decision, can have subconscious effects on human information processing (see, e.g., Tversky and Kahneman 1974; Nisbett et al. 1981; Hackenbrack 1992; Glover 1997; Hoffman and Patton 1997; Frederickson and Miller 2004). Thus, even if users do not perceive the disavowal to be informative, its mere presence could reduce their reliability assessments and affect their investment decisions.

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[10] This holds true in our study; users' reliability assessments are negatively correlated with their investment decisions ($p < 0.02$).

On the other hand, users may view a disavowal as a strategic, self-serving disclosure, and thus not credible. Management generally has incentives to convince users to ignore stock option expense when assessing firm performance, which may lead users to view the disavowal as incentive-consistent. As discussed earlier, attribution and persuasion theory (Kelley 1972; Eagly and Chaiken 1975; Fiske and Pavelchak 1986) suggests individuals consider incentive-consistent signals to be strategic and biased, and thus less credible (Hirst et al. 1995; Hodge et al. 2006). Given that the effect of a stock option disavowal is unclear, we examine this issue on an exploratory basis.[11]

IV. EXPERIMENT

Participants and Design

Our participants are alumni of a major U.S. business school who graduated with a degree in accounting or finance and live within the continental U.S. We selected this participant pool for two reasons. First, SFAC No. 2 indicates that financial statements are intended for users who have "a reasonable understanding of business and economic activities" (FASB 1980, para. 34). Our participant pool meets this criterion, implying that as a group, they have sufficient knowledge to make informed judgments and decisions in our setting. Second, our participants are not professional analysts. The SEC and others have recognized the important role that nonprofessional users of financial statements play in U.S. markets (e.g., see Pitt 2001; Maines and McDaniel 2000; New York Stock Exchange [NYSE] 2001).

We mailed materials to 1,000 individuals selected randomly from the participant pool; 40 packets were returned as undeliverable. Of the remaining 960, we received 220 useable responses, for an overall response rate of 23 percent.[12] On average, our sample consists of relatively sophisticated users of financial statements. Our 220 participants have, on average, 11 years of experience analyzing the financial performance of firms and 16 years of accounting-related work experience. Seventy-nine percent of the participants are Certified Public Accountants.

Our experiment consisted of six cells. We manipulated two variables between participants: (1) reporting environment (voluntary footnote disclosure as allowed under SFAS No. 123, voluntary recognition as allowed under SFAS No. 123, and mandated recognition as required under SFAS No. 123R) and (2) stock option disavowal in the footnotes (presence versus absence).[13]

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[11] Blacconiere et al. (2004) examine the determinants of managements' decision to disavow stock option expense. They provide evidence consistent with disavowals being both informative and strategic. For example, they provide evidence that stock price volatility, an input into option-pricing models, is more difficult to predict for firms that disavow. They also document, however, that disavowing firms have strategic reasons to disavow (e.g., the CEO is receiving excessive compensation).

[12] Our response rate compares favorably with other recent finance and accounting studies that used professionals as participants. For example, Graham et al. (2005) had a response rate of 10 percent using CFOs; Nelson et al. (2002), 16 percent using experienced auditors; Graham and Harvey (2001), 9 percent using CFOs; and Block (1999), 34 percent using Chartered Financial Analysts.

Procedure and Experimental Materials

Each participant received a packet in the mail consisting of three parts: a letter of introduction describing the general purpose of the case and instructions for completing the case, Part 1 of the case, and Part 2 of the case. The instructions for completing the case asked participants to: "Please assume you work in the investment department of a firm. Your supervisor has instructed you to assess the financial performance of Firm Z, which is a member of Industry Q." The instructions also indicated that they would be asked to explain the basis of their assessment.

When making an investment decision, users have prior beliefs about the reliability of financial information. Accordingly, prior to beginning the case, we asked participants to assess the reliability of items commonly found in a firm's financial statements, including stock option compensation expense. For each item (listed in Table 1), we asked participants: "When analyzing firms for the purpose of making investment-related decisions, how reliable do you find the following information?" Participants recorded their answers on an 11-point scale with endpoints labeled "not at all reliable" and "very reliable."

The next part of the case presented background information, which consisted of general information about Firm Z (a hypothetical firm), Firm Z's income statements for 2004, 2003, and 2002, and an abbreviated set of notes (including the stock option footnote). In the disavowal cells, the stock option footnote included a disavowal of the reliability of stock option expense. The wording of the disavowal is identical to Intel's disavowal excerpted in footnote 1. We used two items to manipulate the reporting environment: (1) the placement of stock option expense within the financial report and (2) whether management had a choice to recognize stock option expense. We manipulated these two items as follows.

1. *Placement within financial report*: In all cells, Firm Z's stock option footnote disclosed reported and pro forma net income, as well as reported and pro forma compensation expense, for 2002 through 2004. In the recognition cells, reported net income reflected the effect of stock option expense while pro forma net income did not. In the disclosure cells, reported net income did not reflect the effect of stock option expense while pro forma net income did. Thus, Firm Z's reported net income was lower in the recognition cells. Firm Z's income statements reflected this difference in reported net income.[14]
2. *Recognition choice*: In the disclosure and voluntary recognition cells, Firm Z's stock option footnote stated that SFAS No. 123 gives firms a choice between recognition and disclosure and indicated the approach Firm Z had chosen. In the mandated recognition cells, the stock option footnote stated that the FASB had issued an exposure draft that would require firms to recognize stock option expense (SFAS No. 123R). The note also stated that firms within Firm Z's industry were required, under an SEC pilot study of the exposure draft, to recognize stock option compensation expense.

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[13] We chose not to include a condition where the FASB mandates footnote disclosure because this situation does not correspond to any current, or foreseeable, reporting environment.

[14] Firm Z's income statements did not report stock option compensation expense as a separate line item. Instead, consistent with practice, Firm Z allocated stock option expense across the various expense line items on its income statement. Recognizing stock option expense reduced Firm Z's reported net income by 14 percent.

After reviewing the background material, participants assessed the reliability of stock option earnings relative to non-stock option earnings by answering the following question in the disclosure (recognition) condition:

> Recall that the Financial Accounting Standards Board (FASB) defines *reliable* information as information that is verifiable, reflects a business' activities in an unbiased manner, and is measured with little uncertainty.
>
> Refer to Firm Z's stock option footnote. For 2004, Firm Z disclosed $201 for reported (pro forma) net income and $173 for pro forma (reported) net income. In my opinion, Firm Z's *pro forma* net income is ____________ its *reported* net income (fill in one circle):

○	○	○	○	○	○	○	○	○	○	○
1	2	3	4	5	6	7	8	9	10	11
considerably less reliable than					equally reliable to					considerably more reliable than

Using a question that requires a relative reliability assessment controls for individual differences in participant beliefs about the reliability of financial accounting information in general. Recall that in the recognition (disclosure) conditions, reported (pro forma) net income reflected the effect of stock options expense while pro forma (reported) net income did not. We reverse-coded responses in the recognition conditions to make the responses comparable to those in the disclosure conditions. Our reliability measure, therefore, is a user's perception of stock option earnings relative to non-stock option earnings. A *higher* reliability measure indicates that users perceived stock option earnings—and thus stock option expense—as *more* reliable. Participants also assessed several other items, including the level of Firm Z's "true" 2004 compensation expense. Participants were free to refer to the background material while completing this part of the case. After completing Part 1, participants were instructed to place both the background material and the questionnaire in a marked envelope and seal it.

In Part 2 of the case, we told participants:

> Your company has decided to invest $100,000 in Industry Q. Preliminary analysis has limited the investment options to Firm A, Firm B, and Firm Z (the firm you just analyzed). All three firms are considered industry leaders in terms of market share and innovation. Your supervisor has asked you to decide how much of the $100,000 to invest in each of the three firms.

Participants received background information indicating that the three firms were similar with respect to size, stock price relative to 52-week highs and lows, and history of receiving unqualified audit opinions. Participants also received a ratio described as "critical for evaluating companies in Industry Q" on both a reported and pro forma basis for 2002 through 2004. The ratio was labeled the "human capital productivity ratio" (total sales/total compensation expense). As shown below, we computed this ratio so that Firm Z underperformed the other two firms when total compensation expense included stock option compensation, but outperformed the other two firms when it did not.[15]

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[15] The information presented is for the disclosure condition. The information was identical in the recognition condition except that we reversed the labels "Reported" and "Pro forma."

Human Capital Productivity Ratios

	2004		2003		2002	
	Reported	**Pro Forma**	**Reported**	**Pro Forma**	**Reported**	**Pro Forma**
	Excluding Effect of Stock Options	**Including Effect of Stock Options**	**Excluding Effect of Stock Options**	**Including Effect of Stock Options**	**Excluding Effect of Stock Options**	**Including Effect of Stock Options**
Firm A	13.89	11.15	13.81	10.79	13.01	10.37
Firm B	13.96	11.02	13.68	10.84	13.04	10.41
Firm Z	14.07	10.61	13.99	10.36	13.24	9.97
Industry Q	13.12	9.41	12.83	9.22	12.58	8.88

Changing the relative rank of Firm Z was necessary to create the opportunity for the stock option disavowal to affect participant investment decisions. If the disavowal causes users to discount stock option compensation expense, they should favor Firm Z relative to those not exposed to the disavowal. Participants made their investment decision by indicating the percentage of $100,000 (from 0 percent to 100 percent) they chose to invest in Firm Z.

Participants next answered two manipulation check questions, two questions related to a within-subjects disavowal manipulation, and questions designed to capture demographic information. We discuss these items more fully in the "Results" section. Upon completion of the demographic questions, we instructed participants to insert all materials into the preaddressed, postage-paid return envelope and put it in the mail.

V. RESULTS

Manipulation Checks

The post-experiment questionnaire contained two manipulation check questions. The first pertained to the reporting environment manipulation; the second pertained to the disavowal manipulation. The reporting environment question asked participants (1) whether Firm Z had a choice to recognize or disclose stock option expense and (2) whether Firm Z chose recognition or disclosure. The disavowal question asked participants whether Firm Z disavowed the stock option expense. A total of 174 participants correctly answered both manipulation check questions. The results below are based on the responses of these 174 participants.[16]

Ex Ante Reliability Beliefs

Recall that before beginning the case, participants assessed the reliability of 14 items commonly found in a firm's financial statements. Table 1 lists the items, along with their mean, median, minimum, and maximum reliability ratings. We also report in Table 1 three composite scores: "financial statements," "annual report," and "all." The financial statements composite measure is the simple average of participant assessments of three items: the income statement, balance sheet, and cash flow statement. The annual report composite measure is the simple average of participant assessments of four items: the income statement, balance sheet, cash flow statement, and footnotes. The "all" composite measure is the simple average of all 14 items.

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[16] By including in our tests only those participants who correctly answered our manipulation check questions (i.e., those that acquired the information), we are able to rule out the possibility that our results are due to differences in information acquisition across conditions.

TABLE 1
Participants' *Ex Ante* Reliability Assessments[a]

Item	n	Mean	Median	Minimum	Maximum
The income statement	153	8.14	8.00	3.00	11.00
Revenues	171	8.19	8.00	2.00	11.00
Interest	169	8.47	9.00	1.00	11.00
Net income	167	7.57	8.00	2.00	11.00
The balance sheet	134	8.13	8.00	1.00	11.00
Accounts receivable	173	7.98	8.00	3.00	11.00
Inventory	173	7.53	8.00	4.00	11.00
Property, plant and equipment	172	7.81	8.00	3.00	11.00
Intangible assets	171	5.65	6.00	1.00	11.00
The cash flow statement	166	8.33	9.00	1.00	11.00
The footnotes	130	7.88	8.00	3.00	11.00
Stock option compensation	173	6.49	7.00	1.00	11.00
Research and development	173	6.75	7.00	1.00	11.00
Pensions	173	6.72	7.00	1.00	11.00
Composite measures[b]					
Financial statements	172	8.22	8.42	4.00	11.00
Annual report	172	8.14	8.29	4.00	11.00
All	173	7.51	7.57	3.79	10.69

[a] We asked participants, "When analyzing firms for the purpose of making investment-related decisions, how reliable do you find the following information?" Participants recorded their responses on 11-point scales with endpoints labeled "not at all reliable" and "very reliable."

[b] Our "financial statements" composite measure is the simple average of a participant's assessment of the income statement, balance sheet, and cash flow statement. Our "annual report" composite measure is the simple average of a participant's assessment of the income statement, balance sheet, cash flow statement, and footnotes. Our "all" composite measure is the simple average of all of a participant's assessments. We averaged participants' reliability assessments over the number of individual items they rated. Thus, a participant who failed to assess an individual item, but who assessed other items in a given composite measure, is included in the composite summary statistics reported above. This methodology results in the sample sizes of our composite measures exceeding the sample sizes of some individual measures.

No significant differences exist across the six cells for any of the 14 individual items (p-values range from 0.12 for net income to 1.00 for pensions, two-tailed), including stock option compensation ($p = 0.80$, two-tailed). These results suggest that random assignment of participants to conditions was effective. Several results deserve mention: (1) participants view financial statements as more reliable than footnotes, consistent with H1 (8.22 versus 7.88; $p < 0.01$ two-tailed); (2) participants perceive net income as the least reliable of the income statement items ($p < 0.01$, two-tailed); (3) participants perceive intangible assets as the least reliable of the 14 items ($p < 0.01$, two-tailed); and (4) participants perceive stock option compensation as significantly less reliable ($p \leq 0.05$, two-tailed) than all but two (pensions and intangible assets) of the 14 items listed.

Tests of User Reliability Judgments

Panel A of Table 2 reports descriptive statistics for participant reliability assessments. H1 predicts that user reliability assessments in the mandated recognition condition will exceed their reliability assessments in the voluntary recognition condition, which in turn will exceed their reliability assessments in the footnote disclosure condition. The means across these three reporting environments are 6.47, 5.75, and 4.94, respectively. The Jonckheere-Terpstra test for ordered alternatives reported in Panel B of Table 2 is significant

TABLE 2
Descriptive Statistics and Analysis of Reliability Assessments[a]

Panel A: Mean [Standard Deviation] for Reliability Assessments by Reporting Environment and Disavowal

		Reporting Environment		
	Disclosure	Voluntary Recognition	Mandated Recognition	Overall
Disavowal				
Presence	4.88 [2.85] n = 26	5.74 [3.17] n = 27	6.64 [2.66] n = 25	5.74 [2.96] n = 78
Absence	4.98 [2.35] n = 41	5.76 [2.89] n = 29	6.31 [2.56] n = 26	5.57 [2.61] n = 96
Overall	4.94 [2.53] n = 67	5.75 [3.00] n = 56	6.47 [2.59] n = 51	

Panel B: Jonckheere-Terpstra Test for Ordered Alternatives

Test	J	n	p-value[b]
Is 6.47 > 5.75 > 4.94?	2.96	174	0.002

Panel C: ANOVA for Reliability Assessments

	Sum of Squares	df	Mean Square	F	p-value[c]
Intercept	5530.11	1	5530.11	742.40	0.000
Reporting environment	68.12	2	34.06	4.57	0.012
Disavowal	0.24	1	0.24	0.03	0.859
Environment × Disavowal	1.39	2	0.70	0.09	0.911
Error	1251.42	168	7.45		
Total	6875.00	174			

Panel D: Planned Comparisons to Test Effect of Reporting Environment on Reliability Assessments

Planned Comparison	df	t-statistic	p-value[b]
Mandated recognition > Voluntary recognition	168	1.37	0.086
Mandated recognition > Disclosure	168	3.01	0.002
Voluntary recognition > Disclosure	168	1.64	0.052

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[a] Our reliability measure captures a user's perception of the reliability of the net income number that includes stock option compensation expense (stock option earnings) relative to the reliability of the net income number that excludes stock option compensation expense (non-stock option earnings).

[b] One-tailed.

[c] Two-tailed.

(J = 2.96; $p < 0.01$, one-tailed), as is the reporting environment main effect from the ANOVA results reported in Panel C of Table 2 ($p < 0.01$, one-tailed). These results support H1.

In Panel D of Table 2 we report planned comparisons used to test the individual comparisons implicit in H1. Results show that users perceive mandated recognition to be more reliable than both voluntary recognition ($p < 0.09$, one-tailed) and footnote disclosure ($p < 0.01$, one-tailed). Results also reflect that users perceive voluntary recognition to be more reliable than footnote disclosure ($p < 0.06$, one-tailed). These results are consistent with our ordered alternative and ANOVA results and support H1.

Tests of User Investment Amount

In this section we test whether mandated recognition, voluntary recognition, and voluntary footnote disclosure affect user investment amounts. Hypothesis 2 predicts that the amount users invest in Firm Z in the voluntary recognition condition will exceed the amount they invest in the mandated recognition condition, which in turn will exceed the amount they invest in the footnote disclosure condition. As shown in Panel A of Table 3, descriptive statistics for investment amount are consistent with the ordering predicted in H2. Of the $100,000 available to invest, users invested 43 percent in Firm Z in the voluntary recognition condition, 39 percent in the mandated recognition condition, and 36 percent in the footnote disclosure condition.

As discussed earlier, it is necessary to control for user reliability assessments when testing H2. Accordingly, we test H2 using ANCOVA and planned comparisons, with user reliability assessments of stock option expense as a covariate. The reporting environment main effect from the ANCOVA reported in Panel B of Table 3 is significant ($p < 0.09$, one-tailed), consistent with H2, but the results of the planned comparisons reported in Panel C are mixed. They show that users in the voluntary recognition condition invested more in Firm Z than users in the footnote disclosure condition (43 percent versus 36 percent; $p < 0.04$, one-tailed), which supports H2, but fail to show that they invested more than users in the mandated recognition condition (43 percent versus 39 percent; $p > 0.31$, one-tailed). They also fail to show that users in the mandated recognition condition invested more than users in the voluntary footnote disclosure condition, though the comparison approaches significance (39 percent versus 36 percent; $p < 0.11$, one-tailed).[17]

The Effects of Disavowals

As indicated earlier, we examine the effect of disavowals on user reliability assessments and investment amounts on an exploratory basis. The results reported in Panel C of Table 2 indicate that the presence of the disavowal did not have a main effect on reliability assessments, nor did it moderate the effect of reporting environment on reliability assessments. Similarly, the results reported in Panel B of Table 3 indicate that the presence of the disavowal did not have a main effect on the investment amount, nor did it moderate the effect of reporting environment on the investment amount. These results suggest that disavowals do not influence user reliability judgments or investment amounts.

However, other results from our experiment provide some evidence that disavowals affect some user judgments. In Part 1 of the case, participants assessed Firm Z's "true"

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[17] Unlike H1, we do not use an ordered alternatives test to evaluate H2 because such tests do not allow for covariates. However, running the Jonckheere-Terpstra test for ordered alternatives using investment amount as the dependent variable shows a significant ordered effect ($p < 0.08$, one-tailed), consistent with H2.

TABLE 3
Descriptive Statistics and Analysis of Investment Amount[a]

Panel A: Mean [Standard Deviation] for Investment Amount by Reporting Environment and Disavowal

	Disclosure	Reporting Environment: Voluntary Recognition	Reporting Environment: Mandated Recognition	Overall
Disavowal				
Presence	37% [23%] n = 25	42% [30%] n = 27	42% [26%] n = 25	41% [26%] n = 77
Absence	36% [24%] n = 41	44% [26%] n = 29	37% [23%] n = 26	38% [24%] n = 96
Overall	36% [24%] n = 66	43% [27%] n = 56	39% [24%] n = 51	

Panel B: ANCOVA for Investment Amount

	Sum of Squares	df	Mean Square	F	p-value[b]
Intercept	75996.95	1	75996.95	123.13	0.000
Reliability[c]	3955.06	1	3955.06	6.41	0.012
Reporting environment	2144.65	2	1072.32	1.74	0.179
Disavowal	181.32	1	181.32	0.29	0.589
Environment × Disavowal	429.95	2	214.98	0.35	0.706
Error	102457.45	166	617.21		
Total	375439.90	173			

Panel C: Planned Comparisons to Test Effect of Reporting Environment on Investment Amounts

Planned Comparison	df	t-statistic	p-value[d]
Voluntary recognition > Mandated recognition	166	0.49	0.314
Voluntary recognition > Disclosure	166	1.81	0.036
Mandated recognition > Disclosure	166	1.25	0.107

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[a] We captured participants' investment amount by asking them "What percentage of $100,000 (from 0% to 100%) would you invest in Firm Z?"
[b] Two-tailed.
[c] Participants' reliability assessments are entered as a covariate.
[d] One-tailed.

2004 total compensation expense. Results (not tabled) suggest that exposure to the disavowal weakened user beliefs that Firm Z's compensation expense should include stock option expense. Specifically, exposure to the disavowal decreased the likelihood that participants believed "true" compensation expense should include the cost of stock options (20 percent versus 36 percent; $p < 0.01$, two-tailed) and increased the likelihood that participants believed "true" total compensation expense fell somewhere between the compensation amount that included versus excluded stock option expense (50 percent versus 35 percent; $p < 0.03$, two-tailed).

After answering the manipulation-check questions, participants saw two versions of Firm Z's stock option footnote, side-by-side. The two footnotes were identical except that one included the disavowal and the other did not.[18] Participants then assessed whether the additional disavowal language would affect their reliability assessment of Firms Z's stock option expense and/or the amount they would invest in Firm Z. For the reliability question, only 19 percent of the participants indicated that the disavowal would not affect their reliability assessment (19 percent < 100 percent; $p < 0.01$). Seventy-five percent of the participants indicted that the disavowal would lead them to lower their reliability assessment, which is significantly more than the 6 percent who would raise their assessment ($p < 0.01$) or the 19 percent who would not change their assessment ($p < 0.01$). The effect of the disavowal is weaker for the investment decision, but still significant. Sixty-four percent of the participants indicated that the disavowal would not affect their investment decision (64 percent < 100 percent; $p < 0.01$). Twenty-four percent indicated that the disavowal would cause them to invest less in Firm Z, while 12 percent indicated that the disavowal would cause them to invest more (24 percent > 12 percent; $p < 0.01$).[19] These results suggest that in settings where users are comparing two firms side-by-side—one that disavows and one that does not—disavowals may affect user judgments and decisions.

VI. CONCLUSION

Much controversy exists over the reliability of stock option expense and the consequences of changing from SFAS No. 123 to SFAS No. 123R. The evolution of stock option accounting from the initial adoption of SFAS No. 123 to the passage of SFAS No. 123R encompasses three reporting environments for stock option expense: (1) voluntary footnote disclosure, (2) voluntary income statement recognition, and (3) mandated income statement recognition. In this study we report the results of an experiment designed to examine how these three reporting environments affect relatively sophisticated financial statement user assessments of the reliability of stock option expense and the amount they would invest in the reporting firm.

Results indicate that mandating income statement recognition, as under SFAS No. 123R, increases user assessments of the reliability of stock option expense over the reporting options available under SFAS No. 123 in two ways. First, a mandate from an independent and well-informed authoritative body (the FASB) signals to users a level of reliability exceeding the level signaled by a voluntary, and thereby possibly biased, management choice. Second, voluntary recognition, because it is viewed as incentive inconsistent (i.e., reduces reported net income), signals a level of reliability exceeding the level signaled by voluntary footnote disclosure. With respect to investment amounts, we found that users were more likely to invest in the firm when it voluntarily chose income statement recognition over footnote disclosure, even though this choice reduced reported net income. Finally, our experiment provides evidence about whether an explicit management disavowal of stock option expense affects user reliability assessments and investment decisions. We found that the results reported above were insensitive to whether management disavowed the reliability of stock option expense, but we do provide some evidence in supplemental tests that disavowals have the potential to influence users.

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[18] One footnote was identical to the footnote participants had seen earlier and the other differed only by the presence or absence of the disavowal.

[19] Compared to participants exposed to the disavowal in the between-subjects manipulation, participants not exposed to the disavowal responded no differently to the within-subject reliability or investment choice questions (all p-values > 0.60, two-tailed).

When deciding the appropriate accounting treatment for an item, accounting policymakers face several decisions. They must decide, for example, whether the item is most appropriately disclosed in the footnotes or recognized on the face of a financial statement, and they must decide how much discretion to give management in making this choice. In the context of stock option expense, we demonstrate that when management is given discretion, and they choose income statement recognition over footnote disclosure, users judge the expense to be more reliable and invest more in the firm. When the FASB mandates income statement recognition, users further boost their reliability assessments. These findings are important because they identify important consequences associated with two distinct elements implicit in many accounting policy decisions in the context of an important and highly controversial accounting issue: stock options. In addition, these findings are robust in that they appear to be unaffected by management's attempt to discredit the reliability of the stock option expense by disavowing it in the footnotes.

Our finding that users assess stock option expense voluntarily recognized by management to be more reliable than stock option expense disclosed in the footnotes also provides evidence that reliability issues may partially explain why users treat recognized and disclosed items differently. Prior research has conjectured that voluntary recognition of an item may be informative about the reliability of the item—and thus affect the use of the item, but archival research attempting to examine this issue faces design and/or econometric issues that limit its ability to provide conclusive evidence. Our use of an experiment allowed us to directly assess user reliability assessments while controlling for alternative explanations.

With respect to disavowals, our between-subject manipulation failed to produce an effect, while our within-subject manipulation showed that some users believe that disavowals would affect their judgments and, to a lesser extent, their decisions. These disparate results suggest that further research examining the conditions under which disavowals influence user judgments and decisions may be warranted. Moreover, the possibility that disavowals may be effective under certain conditions suggests that policymakers should monitor them since effective disavowals may serve to undermine the standard-setting process.

Our study is subject to several limitations. First, we provided participants with only a subset of the information typically available to most investors so that they could complete the task in a reasonable amount of time. Second, we mailed the experimental instrument to participants and cannot be certain that they complied with the instructions, introducing the possibility that uncontrolled, extraneous factors affected the results. Notwithstanding these limitations, our results provide important information to managers, accountants, and standard setters about the consequences of initiating standards that eliminate choice and mandate placement, as well as introduce standard setters to disavowing behavior and its potential implications.

REFERENCES

Aboody, D., M. Barth, and R. Kasznik. 2004. Firms' voluntary recognition of stock-based compensation expense. *Journal of Accounting Research* (May): 123–150.

Accounting Principles Board (APB). 1972. *Accounting for Stock Issued to Employees*. Opinion No. 25. New York, NY: America Institute of Certified Public Accountants.

Bagnoli, M., and S. Watts. 2005. Conservative accounting choices. *Management Science* 51: 786–801.

Ball, R., and P. Brown. 1968. An empirical evaluation of accounting income numbers. *Journal of Accounting Research* (Autumn): 159–178.

Barth, M., G. Clinch, and T. Shibano. 2003. Market effects of recognition and disclosure. *Journal of Accounting Research* (September): 581–609.

Barton, J., and G. Waymire. 2004. Investor protection under unregulated financial reporting. *Journal of Accounting and Economics* 38: 65–116.

Beaver, W. 1968. The information content of annual earnings announcements. *Journal of Accounting Research* (Supplement): 67–92.

Bernard, V., and K. Schipper. 1994. Recognition and disclosure in financial reporting. Working paper, University of Michigan.

Blacconiere, W., J. Frederickson, M. Johnson, and M. Lewis. 2004. Voluntary disclosures that disavow mandated disclosures: The case of stock options. Working paper, Indiana University.

Block, S. 1999. A study of financial analysts: Practice and theory. *Financial Analysts Journal* (July/August): 86–95.

Botosan, C., and M. Plumlee. 2001. Stock option expense: The sword of Damocles revealed. *Accounting Horizons 15* (December): 311–328.

Cotter, J., and I. Zimmer. 2003. Disclosure versus recognition: The case of asset revaluations. *Asia-Pacific Journal of Accounting and Economics* 10: 81–99.

Crooch, M. 2005. Presentation at the Financial Reporting Conference. Seattle, Washington, May.

Davis-Friday, P., L. Folami, C. Liu, and H. Mittelstaedt. 1999. The value relevance of financial statement recognition versus disclosure: Evidence from SFAS No. 106. *The Accounting Review* (October): 403–423.

———, C. Liu, and H. Mittelstaedt. 2004. Recognition and disclosure reliability: Evidence from SFAS No. 106. *Contemporary Accounting Research* (Summer): 399–429.

Eagly, A., and S. Chaiken. 1975. An attribution analysis of the effect of communicator characteristics on opinion change: The case of communicator attractiveness. *Journal of Personality and Social Psychology* (July): 136–144.

Elayan, F., K. Pukthuanthong, and R. Roll. 2005. Investors like firms that expense employee stock options and they dislike firms that fail to expense. *Journal of Investment Management* 3 (First Quarter).

Espahbodi, H., P. Espahbodi, Z. Rezaee, and H. Rehranian. 2002. Stock price reaction and value relevance of recognition versus disclosure: The case of stock-based compensation. *Journal of Accounting and Economics* 33 (3): 343–373.

Financial Accounting Standards Board (FASB). 1980. *Qualitative Characteristics of Accounting Information.* Statement of Financial Accounting Concepts No. 2. Norwalk, CT: FASB.

———. 1984. *Recognition and Measurement in Financial Statements of Business Enterprises.* Statement of Financial Accounting Concepts No. 5. Norwalk, CT: FASB.

———. 1995. *Accounting for Stock-Based Compensation.* Statement of Financial Accounting Standards No. 123. Norwalk, CT: FASB.

———. 2004. *Shared Based Payments.* Statement of Financial Accounting Standards No. 123R. Norwalk, CT: FASB.

Fiske, S., and M. Pavelchak. 1986. Category-based versus piecemeal-based affective responses: Developments in schema-triggered affect. In *Handbook of Motivation and Cognition,* edited by R. Sorrentino, and E. Higgins, 167–203. New York, NY: Guilford Press.

Frederickson, J., and J. Miller. 2004. The effects of pro forma earnings disclosures on analysts' and nonprofessional investors' equity valuation judgments. *The Accounting Review* 79 (July): 667–686.

Frost, C. 1997. Disclosure policy choices of UK firms receiving modified audit reports. *Journal of Accounting and Economics* 23: 163–187.

Fuller, R., and R. Metcalf. 1978. Management disclosures: Analysts prefer facts to management's predictions. *Financial Analysts Journal* (March–April): 55–57.

Glover, S. 1997. The influence of time pressure and accountability on auditor's processing of nondiagnostic information. *Journal of Accounting Research* 35 (Autumn): 213–226.

Graham, J., and C. Harvey. 2001. The theory and practice of corporate finance: Evidence from the field. *Journal of Financial Economics* (May/June): 187–243.

———, ———, and S. Rajgopal. 2005. The economic implications of corporate financial reporting. *Journal of Accounting and Economics* 40 (December): 3–73.

Hackenbrack, R. 1992. Implications of seemingly irrelevant evidence in audit judgments. *Journal of Accounting Research* (Spring): 126–136.

Harper, R., W. Mister, and J. Strawser. 1987. The impact of new pension disclosure rules on perceptions of debt. *Journal of Accounting Research* 25 (Autumn): 327–330.

Hirst, E., L. Koonce, and P. Simko. 1995. Investor reactions to financial analysts' research reports. *Journal of Accounting Research* (Autumn): 335–351.

———, and P. Hopkins. 1998. Comprehensive income reporting and analysts' valuation judgments. *Journal of Accounting Research* 36 (Supplement): 47–75.

———, ———, and J. Wahlen. 2004. Fair values, income measurement, and bank analysts' risk and valuation judgments. *The Accounting Review* (April): 453–472.

Hodge, F. 2003. Investors' perceptions of earnings quality, auditor independence, and the usefulness of audited financial information. *Accounting Horizons* (Supplement): 37–48.

———, S. Kennedy, and L. Maines. 2004. Does search-facilitating technology improve the transparency of financial reporting? *The Accounting Review* 79 (3): 687–703.

———, P. Hopkins, and J. Pratt. 2006. Management reporting incentives and classification credibility: The effects of reporting discretion and reputation. *Accounting, Organizations and Society* 31 (October): 623–634.

Hoffman, V., and J. Patton. 1997. Accountability, the dilution effect, and conservatism in auditor's fraud judgments. *Journal of Accounting Research 35* (Autumn): 227–237.

Hopkins, P. 1996. The effect of financial statement classification of hybrid financial instruments on financial analysts' stock price judgments. *Journal of Accounting Research* (Supplement): 33–50.

Hughes, S. 2005. Second Update: SEC OKs delay in options expensing. *Dow Jones Newswire* (April 14).

Imhoff, E., R. Lipe, and D. Wright. 1993. The effects of recognition versus disclosure on shareholder risk and executive compensation. *Journal of Accounting, Auditing and Finance* 8 (4): 335–368.

Kelley, H. 1972. Attribution in social interaction. In *Attributions: Perceiving the Causes of Behavior*, edited by E. Jones, D. Kanouse, H. Kelley, R. Nisbett, S. Valins, and B. Weiner, 151–174. Morristown, NJ: General Learning Press.

Koch, A. 1999. Financial distress and the credibility of management earnings forecasts. Working paper, Carnegie Mellon University.

Levine, C., and J. Hughes. 2005. Management compensation and earnings-based covenants as signaling devices in credit markets. *Journal of Corporate Finance* 11: 832–850.

Libby, R., M. Nelson, and J. Hunton. 2005. Recognition v. disclosure and auditor misstatement correction: The cases of stock compensation and leases. Working paper, Cornell University.

Maines, L., and L. McDaniel. 2000. Effects of comprehensive income volatility on nonprofessional investors' judgments: the role of presentation format. *The Accounting Review* 75 (April): 179–207.

Malkiel, R., and W. Baumol. 2002. Stock options keep the economy afloat. *Wall Street Journal* (April 4): 18.

Nelson, M., J. Elliott, and R. Tarpley. 2002. Evidence from auditors about managers' and auditors' earnings management decisions. *The Accounting Review* (Supplement): 175–202.

Nichols, D., and J. Wahlen. 2004. How do earnings numbers relate to stock returns? A review of classic accounting research with updated evidence. *Accounting Horizons* 18 (December): 263–286.

Nisbett, R., H. Zukier, and R. Lemley. 1981. The dilution effect: Nondiagnostic information weakens the implications of diagnostic information. *Cognitive Psychology* 13 (April): 248–277.

New York Stock Exchange (NYSE). 2001. *Fact Book.* New York, NY: NYSE.

Pitt, H. 2001. Remarks delivered to the AICPA governing council. Miami Beach, Florida, October 22.

Revsine, L., D. Collins, and B. Johnson. 2002. *Financial Reporting and Analysis.* Upper Saddle River, NJ: Prentice Hall.

Securities and Exchange Commission (SEC). 2005. Commission amends compliance dates for FASB statement No. 123R on employee stock options. Press Release 2005-57. April 14. Washington, D.C.: Government Printing Office.

Spence, A. 1974. *Market Signaling: Informational Transfer in Hiring and Related Screening Processes.* Cambridge, MA: Harvard University Press.

Titman, S., and B. Trueman. 1986. Information quality and the valuation of new issues. *Journal of Accounting and Economics* 8 (2): 159–172.

Tversky, A., and D. Kahneman. 1974. Judgment under uncertainty: Heuristics and biases. *Science* 185: 1124–1131.

Wahlen, J. 1994. The nature of information in commercial bank loan loss disclosures. *The Accounting Review* (July): 455–478.

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Auditor-Provided Tax Services: The Effects of a Changing Regulatory Environment

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ABSTRACT: This study investigates auditor-provided tax services from 2000 to 2002, during which fees paid to auditors for nonaudit services (NAS) were disclosed, but separate disclosure of tax service fees was voluntary. We examine changes in the market for tax NAS in 2002, as Congress debated possible prohibition of these services. Using the Heckman MLE procedure to control for selection bias in tax fee disclosure, we find that a strong pre-2002 positive association between tax fees and higher than expected audit fees weakened significantly in 2002, suggesting that some companies paying high audit fees reduced or terminated auditor-provided tax services in 2002. Our findings also suggest that auditor-provided tax services were reduced among new or short-tenure clients. Controlling for tax complexity, we also find some evidence that early "returns" (i.e., tax rate reductions) to companies from auditor-provided tax services were reduced in 2002. Finally, results of our selection model imply that the decision to voluntarily disclose fees paid to the auditor for tax services is positively associated with tax complexity and auditor tenure, and negatively associated with the proportion of NAS fees to total fees. Overall, our findings imply significant change in auditor-provided tax services prior to 2003, when separate disclosure of auditor-provided tax service fees was mandated.

Keywords: *tax fees; tax consulting; audit fees; auditor independence.*

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I. INTRODUCTION

This study examines the market for auditor-provided tax services during 2000–2002, a period in which Congressional debate on possible prohibition of these services likely contributed to fundamental changes in this market. In general, nonaudit service (NAS) provision by auditors is controversial due to potential positive and negative effects.[1] In particular, joint provision of audit and tax services may be beneficial due to efficiencies achieved by having a single professional services firm serve as both auditor and tax consultant. However, potentially damaging outcomes of joint service provision include compromising audit quality and risking shareholder value.

Our study period represents a crucial time in the audit and tax services industry. Large-scale financial frauds focused the attention of the Securities and Exchange Commission (SEC), and ultimately the U.S. Congress, on the issue of audit quality. One important issue in the ensuing debate was whether provision of NAS, including tax services, by a firm's auditor would be allowed or prohibited. Following extensive discussion regarding joint provision of various NAS, the final version of the Sarbanes-Oxley Act of 2002 (SOX) allowed auditor-provided tax services, but disallowed other specific NAS.

We study three basic issues during the limited period that immediately precedes and includes the Congressional debate over SOX in 2002, which may under-represent the full implementation of changes in service relationships resulting from SOX legislation and subsequent regulations. However, this period does provide an initial look at firms' reactions to expected changes in auditor-provided tax services coincident with the SOX debate. First, we examine whether the association between tax fees and audit fees changed during this period of increased pressure on corporations to separate their audit and tax service providers. Second, we investigate whether the association between tax fees and the length of the client/auditor relationship changed over the same period. This question is important because SOX specifically includes auditor tenure as a factor relevant to auditor independence issues (Section 207). While using separate tax and audit service providers should weaken any association between tax fees and auditor tenure, limited prior research (Beck et al. 1988) finds no association between auditor tenure and tax service fees. Third, we examine whether the association between tax service fees and their benefits (as measured by subsequent decline in tax rates) changed during this period. We expect that increasing scrutiny and impending regulatory change likely altered the product mix in the tax services market in 2002, reducing the association between tax fees and tax savings. In studying these issues, we address a specific type of NAS that is interesting and important due to two unique characteristics. First, in contrast to other NAS (e.g., systems development), tax services may have a direct and immediate impact on client income and cash flows through tax rate reduction.[2] Second, tax services may also affect public policy in terms of their potential impact on tax equity. Thus, our service-specific approach complements other research on

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[1] In this paper, "nonaudit services" (NAS) refer to services, other than the financial statement audit, provided by a company's auditor. "Tax fees" refers to fees paid to auditors for tax services.

[2] Phillips et al. (2003) examine earnings management through deferred tax expense. They document that deferred tax expense is incrementally useful over total accruals and discretionary accruals in detecting earnings management to avoid an earnings decline or loss. Their results suggest that deferred tax expense is one source of earnings management.

NAS fees in general, and on the association of those fees with outcomes such as earnings quality.

To conduct our investigation, we use data on tax fees paid to auditors, disclosed in company proxy statements beginning in 2000. Although the SEC did not mandate separate reporting of tax service fees paid to auditors until 2003, numerous firms voluntarily disclosed this information in their proxies from 2000 to 2002. Because firms selectively disclosed fees paid for auditor-provided tax services we use a maximum likelihood estimation (MLE) approach to the Heckman two-step estimation process (Heckman 1979) to control for potential selection bias. Our results suggest that higher abnormal audit fees were positively associated with higher tax fees early in our study period, but this positive association significantly weakened in 2002. This apparent shift in demand for auditor-provided tax services from 2000 to 2002 suggests a preemptive response by firms to separate audit and tax service provision, based on expectations that the impending SOX provisions related to NAS would prohibit or substantially limit auditor-provided tax services.

On the association between tax fees and the length of the auditor-client relationship, we find that higher tax fees were not associated with longer auditor tenure early in the study period (consistent with prior research), but longer tenure was positively associated with higher tax fees in 2002. Although this finding seems to support concerns about NAS affecting auditor independence for longer tenure clients, the longer relationships are also likely to be those in which the client receives higher value. We find a negative association between tax fees and subsequent changes in tax rates prior to 2002 but the association weakened in 2002, suggesting a reduced return on tax-planning investments. In sum, our results are consistent with a change in the fundamental economics of tax service provision to corporations by their auditors in the period leading up to and including the passage of SOX. Also, we infer from the results of our selection model that firm size (assets) and auditor size (Big 5/4) are associated with the decision to purchase tax services from the auditor, but not with the decision to disclose tax fees paid to the auditor. Factors positively associated with voluntary disclosure of tax fees include tax complexity (e.g., foreign tax payments), auditor tenure, and auditor change; voluntary disclosure is negatively associated with the proportion of NAS fees to total fees.

II. BACKGROUND AND HYPOTHESIS DEVELOPMENT

Background: Auditor-Provided Tax Services

When the assurance market reached saturation levels in the 1970s, audit firms expanded their revenue base by providing consulting services in taxation, systems, and other areas. From the 1970s into the 1990s, audit firms' gross revenues from consulting services exceeded those from assurance services. For example, according to reports provided by Zeff (2003, 269), in 1990, Arthur Andersen generated 44 percent of its gross fees from consulting, 35 percent from assurance, and 21 percent from tax. In 2000, PricewaterhouseCoopers generated 49 percent of gross fees from consulting, 33 percent from assurance, and 18 percent from tax. During this period, tax partners experienced increased pressure to generate revenue by cross-selling tax-related services to audit clients, as well as to develop creative tax strategies.

In response to the major corporate financial failures in late 2001 and early 2002, Congress debated and passed SOX, which imposed greater regulation on audit firms, including restrictions on services provided to clients. To promote auditor independence and increase audit quality, SOX prohibits auditors from performing several specific NAS for financial statement audit clients, and requires audit committee approval of permissible NAS. While tax services were not among the NAS prohibited in the final version of SOX, this possibility was under discussion during 2002. Purcell and Lifson (2003) explain that the originally proposed rules prohibited "formulation of tax strategies (tax shelters) designed to minimize a firm's tax obligations" due to a concern about their impairment of auditor independence. However, because tax services can provide benefits to the taxpayer (i.e., lower cost of capital and increased after-tax earnings), the final rules allow audit firms to provide tax services to audit clients, "except for transactions with no business purpose other than tax avoidance (unless consistent with applicable tax laws)." Thus, impending legislation during our study period provided pressure on public companies and their audit firms to abandon jointly provided audit and tax services.

Debate continues among regulators regarding whether joint provision of tax services impairs auditor independence. The Public Company Accounting Oversight Board (PCAOB) published a briefing paper on this issue (PCAOB 2004), held roundtable discussions, and adopted new rules on July 25, 2005. In the roundtable discussions, opponents of tax service provision argued that as with other NAS, tax service fees can lead to lower audit quality. They argued that large revenues and relatively high margins from tax services might affect auditors' judgments on crucial aspects of financial reporting. In addition, they reasoned that auditors cannot objectively judge the appropriateness of highly engineered tax products sold to audit clients by their own firm (Rankin 2004). Proponents of allowing auditors to perform tax services argued that shared knowledge from performing both services can improve audit quality as well as the firm's tax position. If so, then companies benefit because issues associated with the firm's tax returns and/or audited financial statements are more likely to be identified, and identified earlier in the process.

Prior Research on Nonaudit Services

Prior studies have mainly examined two aspects of NAS: (1) the level of NAS purchased, and (2) the possible effects of NAS provision on financial reporting quality. Motivated by longstanding concern about potential effects of economic bonding on auditor objectivity (e.g., Simunic 1984; Beck et al. 1988), recent studies examining the level of NAS purchased (e.g., Dopuch et al. 2003; Brandon et al. 2004) find that financial market participants recognize auditor independence issues associated with NAS and alter their financial decisions accordingly. One specific issue addressed in prior research is whether NAS are associated with higher audit fees. Such an association implies that clients seek to either obtain the benefit of knowledge spillovers between these services, or to influence auditor judgment. While a number of studies document a positive association between NAS fees and audit fees (e.g., Simunic 1984; Simon 1985; Palmrose 1986; Davis et al. 1993; Bell et al. 2001; Antle et al. 2002; Alam and Baez-Diaz 2005), others report no association (Abdel-khalik 1990; O'Keefe et al. 1994; Whisenant et al. 2003).[3]

The length of the auditor-client relationship might also suggest reduced auditor objectivity. Increasing NAS investment with longer auditor tenure may occur due to the auditor's greater client-specific experience and expertise, and/or because greater access to client management improves the ability to identify and sell services to existing clients. While a concern about possible detrimental effects of auditor tenure on audit quality is noted in SOX, recent research (e.g., Myers et al. 2003) finds that earnings quality is actually higher among longer-tenure clients. Evidence on this issue most relevant to the current study is provided by Beck et al. (1988), who studied the association between NAS fees and auditor tenure in 1978–1979. They conclude that while higher NAS fees were associated with longer tenure for some types of service, there was no association of tax NAS with auditor tenure.

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[3] As noted by Firth (2002), a positive association of audit and NAS fees need not imply auditor independence concerns if this association is due to firm-specific factors that increase the cost of both services. However, if these characteristics are controlled for and the association remains strong, auditor-client economic bonding is suggested.

Research on the association between NAS and indicators of financial reporting quality yields inconsistent findings, with some studies (Frankel et al. 2002; Ferguson et al. 2004; Reynolds et al. 2004) suggesting that NAS provision is associated with lower earnings quality, while others studies find no association between earnings quality (or earnings restatements) and NAS provision (e.g., Ashbaugh et al. 2003; Chung and Kallapur 2003; Raghunandan et al. 2003; Kinney et al. 2004). In contrast, Larcker and Richardson (2004) find that higher audit and NAS fees are associated with *higher* earnings quality, concluding that reputation effects constrain auditors from allowing clients to manage earnings. Thus, recent research is mixed on whether NAS compromise auditor independence.

Research in the Tax Service Context and Hypothesis Development

The research reviewed in the previous section provides mixed evidence regarding factors associated with investment in NAS, and the role of NAS in auditor independence. Further research in the specific context of tax services is motivated by several reasons. First, tax services can provide a direct impact on client income and cash flows, in contrast to less directly measurable impacts of other NAS. Supporting differential effects of tax services, Kinney et al. (2004) find that greater investment in tax NAS is associated with a *lower* incidence of earnings restatements (while finding no association between overall NAS fees and earnings restatements). They suggest that this finding implies a possible net benefit in earnings quality from auditor-provided tax services.[4] Second, studies in the U.S. market generally use pre-2001 data, and thus they do not consider the possible preemptive changes in auditor-provided tax services brought about by the SOX debate. Third, Mills et al. (1998) provide a basis for removing variance in tax service fees associated with client complexity, potentially improving the power of hypothesis tests. This study responds to these motivating features of prior research, examining factors associated with investment in tax NAS, and the association of this investment with subsequent changes in tax rates.

We first consider the association between investment in tax fees and the auditor-client relationship during our study period. Hypothesis 1a considers the association between tax fees and audit fees, testing whether the positive association found by some prior studies (e.g., Bell et al. 2001; Antle et al. 2002) holds for auditor-provided tax services during our study period. Of the studies that separately address this issue in the tax context, Palmrose (1986) finds that tax fees are associated with higher audit fees in 1980–1981, while O'Keefe et al. (1994) find no association between tax services and audit fees or hours. Thus, we expect a positive association in H1a; however, we propose in H1b that the association between tax fees and audit fees weakens in 2002, as the approach of new legislation placed pressure on client management and auditors to loosen ties suggesting compromised auditor independence.

H1a: Tax fees paid to auditors are positively associated with audit fees that are greater than expected given client size and risk characteristics.

H1b: The association of tax fees with audit fees will weaken in 2002.

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[4] However, some concern is raised by Favere-Marchesi (2006), who finds that auditors providing tax services to clients assess a lower risk of fraud than auditors who do not.

We also examine the association between tax fees and the length of the auditor-client relationship. While longer relationships could be associated with threats to auditor independence (as implied in SOX), the only direct research evidence on this issue of which we are aware is Beck et al. (1988), who find no association of relatively high NAS fees with auditor tenure. Thus, we propose the following nondirectional hypotheses:

H2a: Tax fees paid to auditors are associated with the length of the auditor-client relationship.

H2b: The association of tax fees with the length of the auditor-client relationship will change in 2002.

Hypotheses 3a and 3b consider the association between investment in tax NAS and subsequent reductions in tax rates, and changes in this association in 2002, respectively. Mills et al. (1998) find that investment in tax services (performed by auditors, nonauditors, and in-house personnel) among large corporations in 1991 is positively associated with tax complexity and negatively associated with tax rates. Consistent with this research, H3a proposes that companies invest in tax planning to reduce future tax liabilities. Thus, higher tax fees should be associated with declining future tax rates, as the effect of the purchased tax strategies becomes evident. Hypothesis 3b proposes a weakening of this association in 2002, due to increasing attention by regulators and the financial press on auditor independence, audit quality, and questionable tax positions.

H3a: Higher tax fees paid to auditors are associated with reductions in future tax rates.

H3b: The association of tax fees with reductions in future tax rates will weaken in 2002.

III. METHODS

Description of the Sample and Variable Definitions

Sample Derivation and Dependent Variable

Data for this study are obtained from Standard & Poor's Audit Fee Data (2000–2002), Compustat Annual Industrial 2003, and Graham's (1996a, 1996b) estimated marginal tax rates (1980–2003). During our study period, registrants were required to report fees paid to auditors in three categories: audit, audit-related, and other. The latter category comprises tax services as well as other NAS. Many firms provided an additional fee category specifically noting tax fees paid to the auditor. Others provided detail on the composition of the "other" category in notes accompanying the proxy statement disclosure. Our dependent variable, *lnTAXFEE*, is the natural log of tax fees paid to auditors, as reported in these sources.

Table 1 summarizes the derivation of the sample. Table 1 indicates that of the 5,727 firm-years with complete data, tax fees are identified in 2,405 cases (approximately 42 percent). For our analyses, firms providing tax fee information are referred to as "disclosers." "Nondisclosers" are firms that provide no tax fee information but for which information was available on all of the independent variables.

Independent Variables—Auditor Relationship Variables

Hypothesis 1a proposes a positive association between tax fees and audit fees that are greater than expected, given client characteristics. We employ an unexpected audit fee measure as our independent variable for testing this hypothesis, because prior literature

TABLE 1
Derivation of the Sample[a]

	Total Sample		Firms Reporting Tax Fee Information (Disclosers)		Firms Not Reporting Tax Fee Information (Nondisclosers)	
	Firm-Years	**Firms**	**Firm-Years**	**Firms**	**Firm-Years**	**Firms**
Initial Sample Based on Standard & Poor's Audit Fee Data 2000–2002	14,434	6,945	5,684	3,693[b]	8,730	5,475[b]
Sample after Combining with Compustat Annual Industrial 2003[c]	12,890	5,946	3,295	2,386	9,582	3,562
Sample after Combining with Graham's Marginal Tax Rate Data 2000–2003[c]	8,044	3,686	3,540	2,234	4,504	2,804
Adjusting for Change in Marginal Tax Rate[c]	5,727	2,925	2,405	1,631	3,322	2,202

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[a] This table presents the derivation of the sample used in the hypothesis testing models.

[b] The numbers do not sum to the first column because some firms did not contribute tax information for all three years in the database.

[c] Firms must have at least two years of available marginal tax rate information to estimate the change in tax rate variables in the model.

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demonstrates that audit fees are highly associated with client size and with other risk-related client characteristics that also appear as independent variables in our model. Specifically, we estimate an OLS model of the log of audit fees on the log of total assets, the ratio of receivables plus inventory to total assets, leverage, return on assets, number of business segments, year, Big N/Non-Big N, and industry. This model explains 78 percent of the variance in audit fees, which is consistent with prior literature. *UNEXPECTED AUDIT FEE* is actual logged audit fees minus forecasted audit fees scaled by logged actual audit fees.

Hypothesis 2a predicts that the length of the auditor-client relationship is associated with investment in auditor-provided tax services. We employ two measures of relationship length. *TENURE* is a continuous variable, representing the number of years that the current auditor has performed the financial statement audit. Additionally, we test whether firms retaining auditors from prior years will have different levels of investment in auditor-provided tax services, relative to firms switching auditors. The independent variable *AUDITOR CHANGE* equals 1 if the firm switched auditors from the previous year, and 0 otherwise. Hypotheses 1b and 2b, which propose that the association between tax fees and auditor relationship variables weakens in 2002, are tested using interactions of auditor relationship variables with the dichotomous variable *YEAR 2002*.

Hypothesis 3a proposes a negative association between investment in auditor-provided tax services in a given year and subsequent year changes in tax rates. Our preferred measure of tax rates is Graham's estimated after-financing marginal tax rate (*MTR*) based on simulations of firms' future taxable income (Shevlin 1990; Graham 1996a, 1996b).[5] We also test the sensitivity of our results using changes in subsequent year effective tax rates (*ETR*). We rely primarily on *MTR* because *ETR* may be too coarse to capture the effects of tax-minimization strategies in the short term, after taking into account firms' current tax environment and capital structure. We compute *ETR* as: Total Taxes (Compustat item #16) minus Deferred Taxes (Compustat item #50) divided by Pretax Income (Compustat item #170). Hypothesis 3b, which proposes that the association between tax fees and tax rate changes weakens in 2002, is tested through the interaction of tax rate change with *YEAR 2002*.

Control Variables

One advantage of studying tax services as a specific NAS is the ability to control for firm characteristics affecting investment in tax services, to improve the power of our hypothesis tests. We use several variables to measure the tax complexity of our sample firms. First, firms with relatively higher deferred taxes are likely to pay higher fees for advice successful in deferring liabilities to future periods. *DEFTAX* is a continuous variable measuring the relative amount of Deferred Taxes (Compustat item #50) to Total Tax Expense (Compustat item #16). Second, firms involved in mergers or acquisitions are likely to need advice on the tax implications of changes in their firm structure and limitations imposed on effectively using the tax attributes of the merging entities or the acquired firm. *MERGER* is a dichotomous variable equal to 1 if the firm engaged in a merger, acquisition, or other corporate restructuring, and 0 otherwise. Third, firms doing business in foreign jurisdictions likely need the assistance of tax professionals in preparing multiple jurisdiction tax returns as well as determining appropriate timing of repatriation of earnings. *FOREIGN* is equal to 1 if taxes were paid to a foreign jurisdiction (Compustat item #64), and 0 otherwise.

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[5] Plesko (2003) questions the Graham marginal tax rate measure because of its limited distribution characteristics. We acknowledge this possible limitation but note that the Graham measure is used extensively in the literature and is one of the only available sources of these estimates.

Fourth, firms with a net operating loss carryforward are likely to seek assistance in scheduling and valuing the related deferred tax assets. *NOL* is equal to 1 if there is an operating loss carryforward (Compustat item #52 > 0) and no pretax income (Compustat item #170 ≤ 0), and 0 otherwise. Fifth, firms with more business *SEGMENTS* (Compustat Business Segments) are likely to have more complex returns and greater need to engage in tax-planning strategies. Finally, firms that expatriate to tax-haven countries are likely to pay higher tax advisory fees to accomplish and maintain this tax minimization strategy (Cloyd et al. 2003; Desai and Hines 2002; Seida and Wempe 2002). *EXPATRIATE* is a dichotomous variable equal to 1 if Cloyd et al. (2003) identified the firm as having relocated its tax home to an offshore tax haven, and 0 otherwise.

Our model also includes control variables for financial characteristics, including size (*lnASSETS*, the natural log of total assets [Compustat item #6]), *LEVERAGE* (total assets minus stockholder equity, divided by total assets [(Compustat item #6 − #60)/#6]), and *Pretax ROA* (pretax income divided by total assets [Compustat item #170/#6]). *LEVERAGE* controls for all of the possible interest expense deductions generated by the firm's borrowing activity regardless of source. While we anticipate positive coefficients for *lnASSETS* and *Pretax ROA*, the appropriate sign on *LEVERAGE* is unclear. The control variable *YEAR 2002* measures the intercept shift associated with that year. Following normal practice in audit fee models, we control for variation in tax fees due to industry membership, using a series of dichotomous variables based on two-digit SIC codes (*INDUSTRY*). We also include dichotomous variables representing Big 5/4 firms (including Arthur Andersen, which was still providing services during part of our sample period).

Empirical Models

The discussion above suggests the following empirical model:

$$\begin{aligned} \mathit{lnTAXFEE} = {} & \alpha_0 + \alpha_1 \mathit{UNEXPECTED\ AUDIT\ FEE} \\ & + \alpha_2 \mathit{UNEXPECTED\ AUDIT\ FEE\ X\ YEAR\ 2002} + \alpha_3 \mathit{TENURE} \\ & + \alpha_4 \mathit{TENURE\ X\ YEAR\ 2002} + \alpha_5 \mathit{AUDITOR\ CHANGE} \\ & + \alpha_6 \mathit{AUDITOR\ CHANGE\ X\ YEAR\ 2002} + \alpha_7 \mathit{CHANGE\ MTR(ETR)} \\ & + \alpha_8 \mathit{CHANGE\ MTR(ETR)\ X\ YEAR\ 2002} + \alpha_9 \mathit{DEFTAX} \\ & + \alpha_{10} \mathit{MERGER} + \alpha_{11} \mathit{FOREIGN} + \alpha_{12} \mathit{NOL} + \alpha_{13} \mathit{SEGMENTS} \\ & + \alpha_{14} \mathit{EXPATRIATE} + \alpha_{15} \mathit{YEAR\ 2002} + \alpha_{16-20} \mathit{AUDITOR\ IDs} \\ & + \alpha_{21} \mathit{lnASSETS} + \alpha_{22} \mathit{LEVERAGE} + \alpha_{23} \mathit{Pretax\ ROA} \\ & + \alpha_{24-87} \mathit{INDUSTRY} + \mathrm{u}. \end{aligned} \qquad (1)$$

Because of the selective disclosure of auditor-provided tax service fees, we use the MLE approach to Heckman's two-step estimation process (Heckman 1979). The Heckman MLE procedure controls for selection bias through simultaneous estimation of a model explaining selection of firms into the group of firms with nonzero reported tax fees (i.e., selection model), with the empirical model above (outcome model).[6,7] Specification of the selection model in our full sample is complex because tax fee disclosure results from a series of management decisions. First, the firm must choose to purchase tax services from an external provider. Research in the corporate tax context (Dunbar and Phillips 2001) proposes that the decision to outsource tax planning is a function of transaction costs relating to human-asset specificity, proprietary technology, and economies of scale, while the outsourcing of tax compliance is a function of firm size and growth. Thus, greater tax complexity should be associated with purchase decisions, as well as with the level of fees. Second, the firm must choose the auditor as provider. We expect that the auditor relationship affects this decision as well as the level of fees, for the same reasons noted above: a close relationship between auditors and client management provides knowledge of client needs and greater access to decision-makers, enabling more sales of services. Thus, the selection model includes tax complexity and auditor relationship variables.

The third decision is that the firm must choose to voluntarily disclose tax fees paid to auditors. While we are aware of no research in the specific context of tax information disclosure, Healy and Palepu (2001) note that in disclosure decisions, managers trade off communication of knowledge about operations against the need to manage reported performance for contracting, political, and governance purposes. Consistent with the disclosure literature, our selection model includes the natural log of assets (*lnASSETS*) and leverage (*LEVERAGE*) as agency variables, as well as pretax ROA (*Pretax ROA*). Additionally, the political sensitivity of some aspects of tax complexity (e.g., foreign taxes and expatriate status) and close auditor relationships might negatively affect disclosure decisions. In addition to tax complexity variables, we include in the selection model two measures of the relative level of NAS fees (*NAS/TOTALFEE*, NAS fees scaled by total fees; and *NAS/*

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[6] The Heckman MLE method tests for correlation between the error terms of the models, and provides an estimate of the inverse Mills ratio and its coefficient (hereafter, the selection bias parameter). We also adjust standard errors for multiple observations on individual firms.

[7] An anonymous reviewer raised two issues related to our test of the association of tax fees and forward changes in tax rates. The first is our choice to use tax fees as the dependent variable, and change in tax rates as the independent variable, in contrast to prior studies that have reversed the position of these variables. There are several reasons why our specification is appropriate in this context. First, the objective of the study is to address multiple hypotheses centered on the level of tax fees paid to auditors. The independent variables in our existing model provide controls for differences in tax fees paid, in order to test these hypotheses. Thus, our current model is the most parsimonious way to achieve the study's multiple objectives. Second, there is a significant selection bias in the voluntary disclosure of tax fees during our sample period. Our current specification provides the ability to use the Heckman procedure to control for this bias.

The second issue raised by the reviewer is the possibility of endogeneity between tax fees and tax rates (in particular, simultaneous endogeneity). Prior related literature varies in perspective regarding whether such a problem exists and, if so, whether it is possible to adequately specify a system of equations to control for it. Selection bias and simultaneity are both endogeneity issues (e.g., Wooldridge 2002). Because selection bias is likely the more serious problem in our analyses, we apply the Heckman MLE approach to control for that specific source of endogeneity. To address possible simultaneity, we use the forward change in tax rates rather than contemporaneous tax rates as used in prior research. However, this may not completely eliminate endogeneity, if it exists.

In additional untabulated analysis, we instrument tax fees using variables drawn from prior research on nonaudit service fees. The R^2 for the instrument equation is 0.71, suggesting that the instrument is reasonably constructed. We then test for endogeneity using the Durbin-Wu-Heckman and Wu-Heckman tests. Results of both tests fail to reject the null hypothesis that our tax fee variable is exogenous. From those results we conclude that, at least in our setting, using change in tax rates and tax fees does not result in a simultaneous endogeneity issue. This supports our use of the Heckman MLE approach to address the selection bias issue, which is of greater concern.

INCOME, NAS fees scaled by pretax income) as additional measures of the political cost of disclosure.[8] The resulting selection model is as follows:

$$\begin{aligned} DISCLOSE = {} & \delta_0 + \delta_1 UNEXPECTED\ AUDIT\ FEE + \delta_2 TENURE \\ & + \delta_3 AUDITOR\ CHANGE + \delta_4 DEFTAX + \delta_5 MERGER \\ & + \delta_6 FOREIGN + \delta_7 NOL + \delta_8 Pretax\ ROA + \delta_9 EXPATRIATE \\ & + \delta_{10} lnASSET + \delta_{11} SEGMENTS + \delta_{12} BIG\ 5/4 \\ & + \delta_{13} LEVERAGE + \delta_{14} NAS/TOTALFEE + \delta_{15} NAS/INCOME \\ & + \delta_{16} YEAR\ 2002 + u. \end{aligned} \tag{2}$$

We view the selection model as preliminary to our primary goal of testing hypotheses on the association of the level of tax fees with other variables. Thus, we provide only a brief discussion of its results in the next section, prior to providing results of our outcome model (Model (1)). As an alternative to the full sample analysis, we apply Heckman MLE to a subsample, containing only disclosers and a smaller sample of nondisclosers that paid the auditor for tax services. Thus, in the subsample model, selection bias is narrowed to the disclosure decision only. This approach conforms to advice by Wooldridge (2002), who notes complications introduced by multiple possible sources of selection bias. To accomplish this, we use the S&P 2003–2004 Audit Fee File to identify firms that paid tax fees to auditors in 2002, but did not report those fees in that year.[9] The subsample contains 300 nondisclosing firms that paid tax fees to their auditor in 2002, combined with 2,405 firm-years disclosing tax fees paid. This procedure specifically addresses the issue of selection bias regarding the decision to disclose fees rather than the combined decision to purchase tax services from the auditor and disclose.

IV. RESULTS

Descriptive Statistics

Table 2 reports descriptive statistics for our sample. Table 2, Panel A, indicates that the mean (median) fee paid by firms reporting auditor-provided tax services approximates $787,000 ($173,000). Comparing tax fees reported by disclosing firms for each year (not tabled), we find that the median fee declines significantly through the sample period ($256,880 in 2000, $191,000 in 2001, and $145,150 in 2002; Pearson Chi-square = 13.14;

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[8] The addition of relative NAS fees to the selection model accomplishes the frequently cited need to include additional variables not contained in the outcome model, in order to avoid multicollinearity. Maddala (1983) and Wooldridge (2002) note that it is not technically necessary for the selection equation to contain additional variables, and caution that adding inappropriate variables to the selection equation can leave the issue of self-selection unaddressed. Stolzenberg and Relles (1997) note that independent variables in the two models are often identical in the sociology literature, and show that the Heckman procedure may not be superior to OLS, even with a large sample size. Based on this caution, we estimate our models several different ways: (1) the Heckman approach with the addition of relative NAS fees, which might affect disclosure but have low correlation with the outcome equation dependent variable; (2) the Heckman approach using only the variables representing tax complexity, auditor relationship, and controls, which we theorize are related to both selection and outcome; (3) censored normal regression, which assumes a random selection process; and (4) OLS regression. While most inferences across the three approaches remain unchanged, control for selection bias using both Heckman models produces some changes in model coefficients, including increased significance of tax complexity. The general similarity of results across these approaches supports the validity of our findings.

[9] Identification is possible because with mandatory disclosure in 2003, firms were also required to report 2002 payments.

TABLE 2
Descriptive Statistics

Variable	Description	Full Sample (n = 5,727)	Firms Reporting Tax Fee Information (Disclosers) (n = 2,405)	Firms Not Reporting Tax Fees (Full Sample Nondisclosers) (n = 3,322)	Firms Not Reporting Tax Fees (Subsample Nondisclosers) (n = 300)
Panel A: Continuous Variables[a]					
TAX FEE	Fee paid to auditor for tax services (thousands)	NA	787 (2,409) 173	NA	NA
AUDIT FEE	Fee for financial statement audit (thousands)	848 (2,110) 285	1,127 (2,824) 334	645 (1,346) 258	800 (1,735) 261
TENURE	Number of years financial statement audit performed by the current auditor	8.98 (9.30) 7	9.92 (9.45) 8	8.30 (9.13) 6	8.50 (8.58) 5
CHANGE MTR	Percentage change in estimated marginal tax rates in the following period (Graham)	−1.41% (12.3) 0%	−1.38% (12.4) 0%	−1.43% (12.3) 0%	−1.86% (13.1) 0%
CHANGE ETR	Percentage change in effective tax rates in the following period	−1.91% (5.92) −0.45%	−2.19% (6.00) −0.94%	−1.85% (5.86) 0.26%	0.86% (6.93) −0.32%
DEFTAX	Current deferred tax expense divided by total tax expense	0.53 (23.4) 0.14	0.45 (18.5) 0.15	0.65 (29) 0.13	0.42 (5.87) 0.12
SEGMENTS	Number of business segments	4.63 (3.16) 4	4.84 (3.29) 4	4.48 (3.06) 4	4.13 (2.58) 4
ASSETS	Total assets (in millions)	6,199 (40,425) 433	9,311 (57,429) 523	3,948 (20,465) 389	2,425 (6,884) 369
LEVERAGE	Total liabilities ÷ Total assets	0.56 (0.44) 0.55	0.56 (0.34) 0.56	0.56 (0.50) 0.55	0.56 (0.36) 0.56
Pretax ROA	Pretax income ÷ Total assets	0.01 (0.76) 0.05	0.01 (0.33) 0.04	0.00 (0.96) 0.05	0.03 (0.26) 0.05

(continued on next page)

TABLE 2 (Continued)

Panel B: Dichotomous Variables[b]

AUDITOR CHANGE	= 1 if changed auditor in the current year, 0 otherwise	9%	14%	6%	9.3%
MERGER	= 1 if merger or acquisition, 0 otherwise	17%	16%	18%	16%
FOREIGN	= 1 if foreign taxes paid, 0 otherwise	42%	46%	39%	31%
NOL	= 1 if a carryforward exists and no pretax income exists, 0 otherwise	15%	14%	15%	13%
EXPATRIATE	= 1 if firm was reincorporated outside the U.S. (Cloyd et al. 2003), 0 otherwise	0.3%	0.4%	0.2%	0.3%

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[a] The table reports the mean (standard deviation) and median for continuous variables.

[b] The table reports the percentage equal to 1 for dichotomous variables. While the percentage changes in tax rates are used here for descriptive purposes, we use decimal changes in the hypothesis testing models.

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p = 0.00). Regarding auditor relationship variables, Table 2 shows that the mean audit fee paid by firms in the full sample is $848,000, and auditor tenure averages about nine years. Overall, the mean of the annual change in *MTR* (*ETR*) is a −1.41 (−1.91) percent. Regarding control variables, the mean ratio of deferred taxes to total taxes is 53 percent, and the mean number of business segments is 4.63. The mean of total assets for the overall sample is about $6.2 billion (median, $433 million). Leverage averages 56 percent and the mean pretax ROA is 1 percent. Table 2, Panel B describes the dichotomous variables. In about 9 percent of the total firm-years, the firm switched auditors. Mergers or acquisitions occurred in 17 percent of the firm-years, taxes were paid to foreign jurisdictions in about 42 percent of sample observations, 15 percent have NOL carryforward amounts, and 0.3 percent have expatriate tax status. Correlations among independent variables (not tabled) are low, with the highest correlation (0.40) between the natural log of assets and the number of business segments.

Selection Model Results, Full Sample, and Subsample

Prior to discussing our outcome model (Model (1)) results, we briefly discuss the results of the selection model (Model (2)) presented in Table 3.[10] The full sample selection model has acceptable discrimination (Hosmer and Lemeshow 2002), as measured by a ROC curve area value of 0.71. The coefficient for the selection bias parameter λ is 0.92, which is significant (p = 0.00). The full sample selection model coefficients that are positive and significant are *TENURE* (p = 0.00); *AUDITOR CHANGE* (p = 0.00), Big 5/4 auditor (p = 0.02); *lnASSETS* (p = 0.00), *FOREIGN* (p = 0.00), and *YEAR 2002* (p = 0.00). These findings are consistent with large, complex entities being more likely to engage in voluntary disclosure in this context. Similar results have been found in other disclosure contexts (e.g., Carcello et al. [2002], for disclosures about audit committees; and Ettredge et al. [2001], for disclosures of financial information on corporate websites). Further, full sample selection model results also indicate that the coefficient for *NAS/TOTALFEE* is negative and significant (p = 0.00). This implies that firms purchasing relatively more NAS services (that were likely exposed to greater political costs during the study period) were less likely to voluntarily disclose separate fees for auditor-provided tax services.

Because there are multiple sources of possible selection bias in the full sample, the subsample selection model may better indicate the factors associated with disclosure, *given that* the firm has decided to purchase tax services from its auditor. The ROC curve area value for the subsample selection model is 0.62, indicating fair discrimination, and the selection bias parameter λ for the subsample is 0.50, which is significant (p = 0.00). The positive and significant coefficients in the subsample selection model are similar to the full sample model, with the exception of the coefficients for *lnASSETS* and *BIG 5/4*, which remain positive but are not significant. These differences between full and subsample are worthy of note, as they imply that firm size and auditor size are positively associated with purchasing auditor-provided tax services (a possible source of bias present only in the full sample), but not the likelihood of disclosing the service once purchased (a possible source of bias present in both samples).

Outcome Model—Full Sample Results

Table 4 presents the results of estimating Model (1) controlling for selection bias. We first discuss the full sample results using *MTR* (Column A), followed by subsample results (Column B), noting any differences arising from use of *ETR* to measure tax spending

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[10] Because there is little guidance in the literature, selection model results are based on two-tailed tests.

TABLE 3
Selection Model Results: Heckman Maximum Likelihood Method

$$\begin{aligned} \mathit{DISCLOSE} = {} & \delta_0 + \delta_1 \mathit{UNEXPECTED\ AUDIT\ FEE} + \delta_2 \mathit{TENURE} \\ & + \delta_3 \mathit{AUDITOR\ CHANGE} + \delta_4 \mathit{DEFTAX} + \delta_5 \mathit{MERGER} \\ & + \delta_6 \mathit{FOREIGN} + \delta_7 \mathit{NOL} + \delta_8 \mathit{Pretax\ ROA} + \delta_9 \mathit{EXPATRIATE} \\ & + \delta_{10} \mathit{lnASSET} + \delta_{11} \mathit{SEGMENTS} + \delta_{12} \mathit{BIG\ 5/4} + \delta_{13} \mathit{LEVERAGE} \\ & + \delta_{14} \mathit{NAS/TOTALFEE} + \delta_{15} \mathit{NAS/INCOME} + \delta_{16} \mathit{YEAR\ 2002} + u \qquad (2) \end{aligned}$$

Variable[a]	A **Full Sample**	B **Subsample**
UNEXP. AUDIT FEE	−0.00[b]	−0.17
TENURE	0.01***	0.01*
AUDITOR CHANGE	0.34***	0.25**
DEFTAX	0.00	0.00
MERGER	0.01	−0.00
FOREIGN	0.18***	0.26**
NOL	0.05	0.08
Pretax ROA	−0.04	−0.09
EXPATRIATE	0.22	0.01
lnASSET	0.08***	0.04
SEGMENTS	0.01	0.02
BIG 5/4	0.18**	0.09
LEVERAGE	−0.10	−0.06
NAS/TOTALFEE	−1.57***	−1.15***
NAS/INCOME	0.00	0.00
YEAR 2002	0.58***	0.07
Constant	−0.63***	1.00***
ROC Curve Area	0.71	0.62
n	5,727	2,705

This table presents results of Heckman MLE selection model, estimating selection into the group of firms disclosing tax fees paid in 2002. Column A presents results of the full sample model, while Column B presents results of the subsample, containing only disclosers and nondisclosers that paid their auditor for tax services.

*, **, *** Indicate significance at 10 percent, 5 percent, and 1 percent, respectively, in two-tailed tests.

[a] Variables are defined in Table 2, with the exception of *NAS/TOTALFEE* (the ratio of nonaudit fees to total auditor fees), and *NAS/INCOME* (the ratio of nonaudit fees to pretax income). Dichotomous variables for industry and for Big 5/4 audit firms are included in the models but are suppressed for efficiency of presentation.

[b] Values reported are the unstandardized regression coefficients.

effectiveness. Results of our hypothesis tests are as follows. Hypothesis 1a predicts that investment in auditor-provided tax services is positively associated with audit fees that are greater than expected in 2000–2001, given client characteristics. The positive and significant coefficient ($p = 0.00$) on *UNEXPECTED AUDIT FEE* supports H1a. The negative and significant coefficient on the interaction of *UNEXPECTED AUDIT FEE* with *YEAR 2002* ($p = 0.02$) supports H1b, which suggests that the positive association between tax-related NAS fees and audit fees weakens in 2002. The marginal effects for these variables, conditional on tax fee reporting, imply a 62.19 percent increase in tax fees for a one standard deviation increase in unexpected audit fees prior to 2002, and a 37.79 percent increase

TABLE 4
Outcome Model Results: Heckman Maximum Likelihood Method

$$
\begin{aligned}
lnTAXFEE = {} & \alpha_0 + \alpha_1 UNEXPECTED\ AUDIT\ FEE + \alpha_2 UNEXPECTED\ AUDIT\ FEE\ X\ YEAR\ 2002 + \alpha_3 TENURE \\
& + \alpha_4 TENURE\ X\ YEAR\ 2002 + \alpha_5 AUDITOR\ CHANGE + \alpha_6 AUDITOR\ CHANGE\ X\ YEAR\ 2002 + \alpha_7 CHANGE\ MTR(ETR) \\
& + \alpha_8 CHANGE\ MTR(ETR)\ X\ YEAR\ 2002 + \alpha_9 DEFTAX + \alpha_{10} MERGER + \alpha_{11} FOREIGN + \alpha_{12} NOL + \alpha_{13} SEGMENTS \\
& + \alpha_{14} EXPATRIATE + \alpha_{15} YEAR\ 2002 + \alpha_{16\text{–}20} AUDITOR\ IDs + \alpha_{21} lnASSETS + \alpha_{22} LEVERAGE + \alpha_{23} Pretax\ ROA \\
& + \alpha_{24\text{–}87} INDUSTRY + u \qquad (1)
\end{aligned}
$$

Variable[a]	Expected Sign	A Full Sample MTR	A Full Sample ETR	B Subsample MTR	B Subsample ETR
Constant		−0.52[b]	−0.51	0.63*	0.64*
UNEXP. AUDIT FEE (H1a)	+	3.38***	3.42***	3.18***	3.21***
UNEXP. AUDIT FEE X YEAR 2002 (H1b)	−	−1.14**	−1.17**	−1.14**	−1.17**
TENURE (H2a)	+/−	0.00	0.00	−0.00	−0.00
TENURE X YEAR 2002 (H2b)	+/−	0.01*	0.01*	0.01*	0.01**
AUDITOR CHANGE (H2a)	+/−	0.38***	0.39***	0.19*	0.19*
AUDITOR CHANGE X YEAR 2002 (H2b)	+/−	−0.39***	−0.40***	−0.34**	−0.35**
CHANGE MTR (ETR) (H3a)	−	−0.33*	−0.09**	−0.43**	−0.09**
CHANGE MTR (ETR) X YEAR 2002 (H3b)	+	0.34	0.06	0.46*	0.06
DEFTAX	+	0.00	0.00	0.00**	0.00***
MERGER	+	0.10*	0.10*	0.12**	0.12**
FOREIGN	+	0.29***	0.29***	0.22***	0.22***
NOL	+	0.06	0.07	0.05	0.06
SEGMENTS	+	0.06***	0.06***	0.06***	0.06***
EXPATRIATE	+	0.58**	0.59**	0.46*	0.47*
YEAR 2002	+/−	0.20***	0.19***	−0.19***	−0.20***
lnASSETS	+	0.55***	0.55***	0.53***	0.53***
LEVERAGE	+/−	0.19**	0.19**	0.23***	0.23***
Pretax ROA	+	0.10*	0.10*	0.12*	0.12*
Pseudo R^2		0.11	0.11	0.16	0.16
n		5,727		2,705	

(*continued on next page*)

TABLE 4 (Continued)

This table presents results of Heckman MLE outcome models, explaining the natural log of fees paid to auditors for tax services as a function of auditor relationship, subsequent year change in marginal tax rates, tax complexity, and control variables. Column A presents results of the full sample models, using marginal or effective tax rates as alternate measures of tax service effectiveness. Column B presents results of the subsample, which contains only disclosers and nondisclosers that paid their auditor for tax services.

*, **, *** Indicate significance at 10 percent, 5 percent, and 1 percent, respectively, in a one-tailed test if directional predictions are made and a two-tailed test otherwise.

[a] Variables are defined in Table 2. Dichotomous variables for industry and for individual Big 5/4 audit firms are included in the models but are suppressed for efficiency of presentation.

[b] Values reported are the unstandardized regression coefficients.

during 2002.[11] The declining association between tax and audit fees suggests that firms were reducing their reliance on the auditor for tax services in 2002.

Hypothesis 2a proposes that the length of the auditor relationship affects investment in auditor-provided tax services, while H2b proposes that this association changes in 2002. The coefficient for *TENURE*, is insignificant; however, the coefficient on the interaction of *TENURE* and *YEAR 2002* is positive and marginally significant ($p = 0.08$). This implies no association prior to 2002 between tax fees and tenure, consistent with Beck et al. (1988). However, this relationship changes in 2002, such that clients with longer auditor relationships had a somewhat higher level of tax fees in 2002 than did those with shorter auditor relationships. Conditional marginal effects imply a 0.61 percent increase per year of tenure and, thus, for example, a ten-year auditor relationship would result in 6.28 percent higher tax fees.

We also test H2a and H2b using *AUDITOR CHANGE* to represent length of the auditor-client relationship. Table 4 shows that the coefficient for *AUDITOR CHANGE* is positive and significant ($p = 0.00$), and the coefficient on the interaction of *AUDITOR CHANGE* with *YEAR 2002* is negative and significant ($p = 0.01$).[12] Conditional marginal effect computations show that firms changing auditors prior to 2002 paid 46.03 percent higher tax fees (ETR model: 47.01 percent), while those changing auditors in 2002 paid 1.41 percent *lower* tax fees (ETR model: 1.00 percent). Taken together with results on *TENURE*, our results suggest that clients with new auditors purchased more tax services prior to 2002. However, in 2002 this positive association weakens, such that clients with short-tenure auditors purchased less tax services from their auditors than clients with long-tenure relationships.

Hypotheses 3a and 3b predict that the purchase of auditor-provided tax services is associated with declining tax rates in the following year, but that this relationship weakens in 2002. The negative and marginally significant coefficient on *CHANGE MTR* ($p = 0.07$), supports H3a. While the positive coefficient for the interaction of *CHANGE MTR* and *YEAR 2002* suggests that the relationship between tax fees and tax rate changes weakens in 2002, the variable is not significant in the full sample model. For the *MTR* model, conditional marginal effects imply that a 1 percent decline in tax rate is associated with 28 percent higher tax fees prior to 2002.

The coefficients for several tax complexity variables are positive and significant, including *MERGER* ($p = 0.08$), *FOREIGN* ($p = 0.00$), *SEGMENTS* ($p = 0.00$), and *EXPATRIATE* ($p = 0.05$). Overall, these findings imply that firms with complex tax characteristics paid higher fees to auditors for tax services, while controlling for client size and other characteristics. For the remaining control variables we find positive and significant coefficients for *lnASSETS* ($p = 0.00$), *LEVERAGE* ($p = 0.03$), and *Pretax ROA* ($p = 0.09$). The *YEAR 2002* intercept shift variable is also positive and significant ($p = 0.01$).

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[11] We evaluate conditional marginal effects at the mean of the distribution. We report the conditional effects (i.e., only for firms reporting tax fees), as the economic impact is more meaningful among firms whose tax service payments are known.

[12] These results are obtained excluding firm-years that involved auditor changes from Arthur Andersen LLP in 2002 due to the possibility that measurement error is associated with the unusual circumstances surrounding the "flight" from Andersen. When the 2002 Arthur Andersen switches are included in the sample, the results for the *AUDITOR CHANGE* and *AUDITOR CHANGE X YEAR 2002* are statistically unchanged although coefficients and p-values are slightly different. However, the 2002 switches from Arthur Andersen to other auditors exhibit greater reduction in tax fees paid to the new auditor than do other auditor switches.

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Outcome Model—Subsample Results

As previously noted, interpretation of our full sample results is complicated by multiple sources of selection bias. This issue motivates re-estimating Model (1) using the subsample of tax service purchasers, shown in Table 4, Column B. Subsample outcome model results for both the *MTR* and *ETR* models are generally consistent with full sample outcome model results with two exceptions. First, the coefficient on the interaction of *CHANGE MTR* and *YEAR 2002* is now negative and marginally significant ($p = 0.10$), providing some support for H3b. Marginal effects in the subsample outcome model indicate 35.14 percent higher fees associated with a 1 percent reduction in tax rate prior to 2002 and 2.64 percent higher tax fees with a 1 percent tax rate increase in 2002. Second, the positive coefficient for *AUDITOR CHANGE* and the negative coefficient on the interaction of *AUDITOR CHANGE* and *YEAR 2002* are consistent with the full sample results but have reduced significance in the subsample outcome model ($p = 0.06$ and $p = 0.02$, respectively). Conditional marginal effects for firms changing auditors prior to 2002 indicate 20.67 percent higher tax fees, while those changing auditors in 2002 paid 14.01 percent lower tax fees. Other differences include the positive coefficient for *DEFTAX*, which is insignificant in the full sample outcome model but is significant in the subsample model ($p = 0.02$), and the coefficient for *YEAR 2002*, which is positive and significant in the full sample outcome model but is negative and significant ($p = 0.01$) in the subsample model.

Supplementary Analyses

We perform two supplementary analyses (not tabled). First, we investigate whether, following the passage of SOX in 2002, investment in auditor-provided tax services continued to decline, or increased once it became clear that these services were not expressly prohibited. Either result is possible, as although regulatory pressure on the industry was maintained during 2003, the final version of SOX did not specifically prohibit selling of tax services by auditors. Using the 2003 S&P Audit Fee File, we matched 2003 tax fee data for 1,124 of the 1,182 sample firms reporting tax fees paid to auditors in 2002. The mean (median) tax fee paid by these firms in 2002 is $709,000 ($148,500), while the mean (median) tax fee reported in 2003 is $646,000 ($127,345), a significant reduction ($p = 0.00$). This suggests that although SOX did not prohibit tax services by auditors, factors such as the need for audit committee review and pressure by regulators continued to depress this market.

Second, we investigate whether investment in auditor-provided tax services is associated with evidence of earnings management during our study period; specifically, whether there is an association between tax fees and discretionary accruals. We perform a preliminary test by first computing abnormal accruals using the modified Jones model (as in, for example, Francis et al. 2005).[13] When we include this measure of earnings management in Model (1), we do not find a significant association between auditor-provided tax services and earnings management, either before or during 2002. This result provides some comfort on the possible effects of high tax service fees on auditor independence. However, it is important to note that our results indicate a negative association between tax fees and tax rates. This implies that firms buying tax services before 2002 obtained a boost in earnings from the decline in tax rate, and thus might have been less motivated to manage earnings in other ways. Further research should investigate this important issue.

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[13] We estimate discretionary accruals using all available firms on the Compustat 2003 file, restricted to industries containing at least ten firms. This approach avoids potential bias in the estimation of accruals using only firms in the target sample.

V. CONCLUSION AND LIMITATIONS

Conclusions

This study examines factors associated with investment in auditor-provided tax services during 2000–2002, a period of change in the regulatory climate for NAS and for tax services specifically. While there has been research directed toward joint audit/NAS provision since disclosure of U.S. audit and NAS fees in 2000, there is little evidence on tax services specifically, despite its importance to companies, providers, and the public interest. This issue is important because while the tax services are not among prohibited NAS, this exemption has not been adequately examined in light of the concerns over auditor independence.[14] Research findings on the association between total NAS and auditor independence are mixed, with some evidence provided by Kinney et al. (2004) that tax services may differ from other types of NAS in this regard. Our research provides three main findings.

First, we find a strong positive association between tax fees and unexpected audit fees during 2000–2001 and a significant reduction in the strength of this positive association in 2002. We also find that that during 2000–2001, new clients paid significantly greater tax fees than continuing clients; however, in 2002, shorter-tenure clients paid lower tax fees. Overall, these findings suggest that the reduction in association between audit and tax service fees in 2002 was due to new and relatively short-tenure clients terminating tax services (or not initiating new services), while services provided by longer-tenure auditors were maintained despite growing pressure by regulators to seek tax services elsewhere. Two prior studies are important in interpreting this result. Kinney et al. (2004) find that higher tax fees paid to auditors are associated with fewer earnings restatements. In addition, Myers et al. (2003) find that longer auditor tenure is associated with higher earnings quality. Combining these results with ours, the evidence suggests that service quality is the main driver of joint audit/tax service provision, rather than attempts by clients to "buy" the auditor's acquiescence. This finding supports the PCAOB's decision not to ban auditors from performing all tax services for their clients, but to limit the prohibition to certain tax services likely to be more problematic. Our supplemental comparison of 2003 tax fees paid by the 2002 reporters in our sample shows a continued decline in this market, even after the final version of SOX did not prohibit auditors from providing tax services to clients.

Our second main finding relates to the association between tax fees and subsequent changes in tax rates. Our results imply that during 2000–2001, higher tax fees were associated with greater reductions in future tax rates. However, we find some evidence that this association weakened during 2002, in which tax fees paid were not associated with future declines in tax rates (controlling for firm size, income, and tax complexity), implying no or reduced "return" to those expenditures. Although we cannot directly observe the nature of tax services provided, our interpretation of these findings is that the services provided in 2002 were less likely to produce reduction in tax rates.

Our third main set of findings is derived from our selection model (Model (2)). The significant selection bias parameter provides evidence that the voluntary disclosure of tax fees during the early 2000s was not a random event. As has been found in other financial disclosure contexts, the disclosure decision varies depending on firm-specific and auditor-specific factors. For instance, the difference in significance of the effects of firm size and auditor size on disclosure between our full sample and subsamples suggests that these factors are positively associated with the decision to *purchase* auditor-provided tax services, but not with the decision to *disclose* them once purchased. Consistent with testimony in the PCAOB hearings on tax services, complex entities need the services of large audit firms (e.g., to prepare the tax provision). Once the auditor is chosen to provide this service, the fee amount is strongly associated with firm size, but the propensity to voluntarily disclose the relationship is unaffected. However, our results also show that (controlling for size and other variables in the model) firms with relatively high ratios of NAS to total fees are less likely to voluntarily disclose the amount of tax fees in the mix. Because the 2000 to 2002 period provides a unique view of voluntary NAS fee disclosure, further research should investigate these issues in greater depth.

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[14] At this writing the PCAOB has proposed further restrictions on the provision of tax services by auditors, which have yet to be ratified by the SEC. See http://www.pcaobus.org/News_and_Events/News/2004-12-14.asp for further discussion of the proposed ban.

Limitations

Our findings are limited by features of our data and method. First, reporting of tax fees paid to auditors was voluntary during our sample period, and disclosure of tax fees paid to nonaudit providers was not required. We attempt to control for this selection bias using the Heckman procedure. While control may not be complete, one of the significant contributions of the current study is to show that during a period of regulatory uncertainty, the provision of tax services by auditors changed from previous patterns. Thus, use of voluntary disclosure data is necessary in order to address our research questions, and this study provides a foundation for further studies of tax fees under mandatory reporting. Second, while we measure tax rate changes over a single year, tax-planning effects are likely to occur over longer horizons. This implies that the observed relationship between change in tax rates and current tax fees may understate the tax benefit of fees paid, especially for effective tax rates. Finally, we observe the tax services market over a limited period, and thus may not fully observe the implementation of changes in service relationships resulting from SOX legislation and subsequent regulations.

REFERENCES

Abdel-khalik, R. 1990. The jointness of audit fees and demand for MAS: A self-selectivity analysis. *Contemporary Accounting Research* 6 (Spring): 295–322.

Alam, P., and A. Baez-Diaz. 2005. The joint provision of audit and non-audit services: Evidence from the U.S. banking industry. Working paper, Kent State University, Kent, OH.

Antle, R., E. A. Gordon, G. Narayanamoorthy, and L. Zhui. 2002. The joint determination of audit fees, NAS fees and abnormal accruals. Working paper, Yale University, New Haven, CT.

Ashbaugh, H., R. LaFond, and B. W. Mayhew. 2003. Do NAS compromise auditor independence? Further evidence. *The Accounting Review* 78 (July): 611–639.

Beck, P. J., T. Frecka, and I. Solomon. 1988. An empirical analysis of the relationship between MAS involvement and auditor tenure: Implications for auditor independence. *Journal of Accounting Literature* 7: 65–83.

Bell, T., W. Landsman, and D. Shackleford. 2001. Auditors' perceived business risk and audit fees: Analysis and evidence. *Journal of Accounting Research* 39: 35–43.

Brandon, D., A. Crabtree, and J. Maher. 2004. Nonaudit fees, auditor independence, and bond ratings. *Auditing: A Journal of Practice & Theory* 23: 2 (September): 89–103.

Carcello, J. V., D. R. Hermanson, and T. L. Neal. 2002. Disclosures in audit committee charters and reports. *Accounting Horizons* 16 (4): 291–304.

Chung, H., and S. Kallapur. 2003. Client importance, nonaudit services, and abnormal accruals. *The Accounting Review* 78 (October): 931–955.

Cloyd, C. B., L. F. Mills, and C. D. Weaver. 2003. Firm valuation effects of the expatriation of U.S. corporations to tax-haven countries. *The Journal of the American Taxation Association* 25 (Supplement): 87–109.

Davis, L., D. Ricchiute, and G. Trompeter. 1993. Audit effort, audit fees, and the provision of nonaudit services to audit clients. *The Accounting Review* 68: 135–150.

Desai, M. A., and J. R. Hines, Jr. 2002. Expectations and expatriations: Tracing the causes and consequences of corporate inversions. *National Tax Journal* 55 (September): 409–440.

Dopuch, N., R. King, and R. Schwartz. 2003. Independence in appearance and fact: An experimental investigation. *Contemporary Accounting Research* 20: 29–114.

Dunbar, A. E., and J. D. Phillips. 2001. The outsourcing of corporate tax function activities. *The Journal of the American Taxation Association* 23: 35–49.

Ettredge, M., V. Richardson, and S. Scholz. 2001. The presentation of financial information at corporate websites. *International Journal of Accounting Information Systems* 2: 149–168.

Favere-Marchesi, M. 2006. The impact of tax services on auditors' fraud risk assessment. *Advances in Accounting* 22: 149–165.

Ferguson, M., G. Seow, and D. Young. 2004. Nonaudit services and earnings management: UK evidence. *Contemporary Accounting Research* 21 (Winter): 613–642.

Firth, M. 2002. Auditor-provided consultancy services and their association with audit fees and audit opinions. *Journal of Business Finance & Accounting* 29 (June/July): 661–693.

Francis, J., R. LaFond, P. Olsson, and K. Schipper. 2005. The market pricing of accruals quality. *Journal of Accounting and Economics* 39 (June): 295–327.

Frankel, R. M., M. F. Johnson, and K. K. Nelson. 2002. The relation between auditors' fees for NAS and earnings management. *The Accounting Review* 77 (October): 71–105.

Graham, J. 1996a. Debt and the marginal tax rate. *Journal of Financial Economics* 41: 41–73.

———. 1996b. Proxies for the corporate marginal tax rate. *Journal of Financial Economics* 42: 187–221.

Healy, P., and K. Palepu. 2001. Information asymmetry, corporate disclosure, and the capital markets: A review of the empirical disclosure literature. *Journal of Accounting and Economics* 31: 405–440.

Heckman, J. 1979. Sample selection bias as a specification error. *Econometrica* 47: 153–161.

Hosmer, D. W., and S. Lemeshow. 2002. *Applied Logistic Regression.* Wiley Series in Probability and Statistics. 2nd edition. New York, NY: John Wiley & Sons, Inc.

Kinney, W., Z-V. Palmrose, and S. Sholtz. 2004. Auditor independence, NAS, and restatements: Was the U.S. government right? *Journal of Accounting Research* 42 (June): 561–588.

Larcker, D. F., and S. A. Richardson. 2004. Fees paid to audit firms, accrual choices, and corporate governance. *Journal of Accounting Research* 42 (June): 625–658.

Maddala, G. S. 1983. *Limited Dependent and Qualitative Variables in Econometrics.* Cambridge, U.K.: Cambridge University Press.

Mills, L., M. M. Erickson, and E. L. Maydew. 1998. Investments in tax planning. *The Journal of the American Taxation Association* 20 (Spring): 1–20.

Myers, J., L. Myers, and T. Omer. 2003. Exploring the term of the auditor-client relationship and the quality of earnings: A case for mandatory auditor rotation? *The Accounting Review* 78 (3): 779–799.

O'Keefe, T., D. Simunic, and M. Stein. 1994. The production of audit services: Evidence from a major public accounting firm. *Journal of Accounting Research* 32 (Autumn): 241–261.

Palmrose, Z-V. 1986. The effect of nonaudit services on the pricing of audit services: Further evidence. *Journal of Accounting Research* 24 (Autumn): 405–411.

Phillips, J., M. Pincus, and S. Olhoft Rego. 2003. Earnings management: New evidence based on deferred tax expense. *The Accounting Review* 78 (2): 491–521.

Plesko, G. A. 2003. An evaluation of alternative measures of corporate tax rates. *Journal of Accounting and Economics* 35 (June): 201–226.

Public Company Accounting Oversight Board (PCAOB). 2004. *Briefing Paper: Auditor Independence and Tax Services Roundtable.* Washington, D.C.: PCAOB.

Purcell, T. J. III, and D. Lifson. 2003. Tax services after Sarbanes-Oxley. *Journal of Accountancy* 196 (November): 32–37.

Raghunandan, K., W. J. Read, and J. S. Whisenant. 2003. Initial evidence on the association between nonaudit fees and restated financial statements. *Accounting Horizons* 17 (September): 223–234.

Rankin, K. 2004. PCAOB reopens tax services debate. *Accounting Today* 18 (14): 1–38.

Reynolds, J. K., G. Deis, and J. Francis. 2004. Professional service fees and auditor objectivity. *Auditing: A Journal of Practice & Theory* 23 (March): 29–52.

Seida, J. A., and W. F. Wempe. 2002. Market reaction to corporate inversion transactions. *Tax Notes* 97 (November): 1098.

Shevlin, T. 1990. Estimating corporate marginal tax rates with asymmetric tax treatment of gains and losses. *The Journal of the American Taxation Association* 11 (1): 51–67.

Simon, D. T. 1985. The audit services market: Additional empirical evidence. *Auditing: A Journal of Practice & Theory* 5 (Fall): 71–78.

Simunic, D. 1984. Auditing, consulting, and auditor independence. *Journal of Accounting Research* 22: 679–702.

Stolzenberg, R., and D. Relles. 1997. Tools for intuition about sample selection bias and its correction. *American Sociological Review* 62 (3): 494–507.

Whisenant, S., S. Sankaraguruswamy, and K. Raghunandan. 2003. Evidence on the joint determination of audit and non-audit fees. *Journal of Accounting Research* 41 (September): 721–744.

Wooldridge, J. 2002. *Econometric Analysis of Cross Section and Panel Data.* Cambridge, MA: The MIT Press.

Zeff, S. A. 2003. How the U.S. accounting profession got where it is today: Part II. *Accounting Horizons* 17 (December): 267–286.

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Do Venture Capitalists Influence the Decision to Manage Earnings in Initial Public Offerings?

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ABSTRACT: Prior studies suggest that venture capitalists (VCs) play a monitoring role. We predict and find that IPO-year abnormal accruals are lower in the presence of VCs for a sample of 2,630 IPO firms during 1983–2001. Our findings are robust to controls for the endogenous choice of VC financing. We consistently find that the VC effect holds even when controlling for IPO lock-up provisions, VC partial cashing out subsequent to the IPO, and alternative proxies for earnings management. In addition, our findings do not support the claims of critics that VCs inflated earnings during the Internet IPO bubble. Finally, we provide some evidence that the lower earnings management associated with VC monitoring partially explains the superior post-IPO returns of VC-backed firms.

Keywords: *earnings management; initial public offerings; venture capitalists; monitoring.*

Data Availability: *Data are publicly available from sources cited in the study.*

I. INTRODUCTION

This paper contributes to several streams of research by examining earnings management in venture capital-backed initial public offerings. To date, accounting research identifies certain parties that constrain management's incentives to manage earnings upward, such as external auditors, boards of directors, audit committees, and actuaries (see, Klein 2002; Gaver and Paterson 2001). Finance theory indicates that venture capitalists (VCs) generally add value to their investee ("portfolio") companies, not only through their

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cash infusions, but also through services such as business plan development, networking, human resource management, and post-investment monitoring and oversight (Gompers and Lerner 2002; Hellman and Puri 2002). Researchers generally substantiate the perceived monitoring role with empirical findings that VC-backed IPOs exhibit relatively less first-day underpricing and higher long-run stock returns (Megginson and Weiss 1991; Brav and Gompers 1997; Gompers and Lerner 2002).

Despite extensive literature and anecdotal evidence regarding VC-added value in the public equity market, little is known empirically about which specific management decisions monitored by VCs actually protect or increase long-term shareholder value; even less is understood to date about whether VCs actually impact financial reporting and disclosure choices of public firms. However, VCs themselves claim that they bring "exacting oversight to management's preparation and the auditor's review of the financial statements" (Heesen 2002). Teoh et al. (1998b) find an inverse relationship between positive IPO-related abnormal accruals and long-run stock performance. Hence, we believe the earnings management decision in an IPO setting provides a powerful context in which to address these open questions. Our study specifically asks whether VCs constrain opportunistic earnings management around IPOs (Dechow and Skinner 2000; Teoh et al. 1998).

Examining the impact of VCs on earnings management in an IPO setting is important for several reasons. First, VCs are significant players in the IPO market. Statistics show that, on average, over the past decade one in three IPOs received VC backing (Gompers and Lerner 2002). Venture capitalists typically require: seats on the board of directors of their portfolio companies, regular interim financial reports, and access to key managers and managers' decisions as conditions for their investment and ongoing monitoring of the firm (Gompers and Lerner 2002). To date, evidence that indicates whether these monitoring activities affect financial reporting and earnings management decisions is sparse.

Second, critics of the VC industry suggest that the interests of VCs conflict with those of the other pre- or post-IPO stakeholders around the time of the offering (Stross 2000; Healy 2002; Tunick 2003; Cohen and Langberg 2005). Their perception is that VCs use their influence over management to artificially inflate IPO firm stock prices and, thus, were partly responsible for fueling the Internet IPO bubble (e.g., see Biddle 2001; Buckman 2001; Mills 2001). Recently, the SEC has proposed regulation for hedge funds (and potentially for the venture capital industry), citing improper accounting and reporting behavior by funds as motivation for the proposed rules (SEC Release No. IA-2266; Borrell 2004). Consequently, providing evidence on whether VCs are associated with management financial reporting behavior is timely and relevant.

We examine a sample of IPO firms in the period 1983–2001. We predict that IPO-year abnormal accruals would be lower in the presence of VCs. The structure of our tests and sensitivity analyses is as follows: (1) we regress abnormal accruals on VC presence, VC incentives to monitor, and appropriate control variables; (2) we perform a battery of tests to ensure that our main results are actually driven by VCs and not by competing explanations (e.g., the endogenous choice of VC financing, lock-ups, VC exits, the Internet IPO bubble, etc.); (3) we confirm that our findings are robust across various earnings management indicators (e.g., cross-sectional Jones and performance-matched abnormal accruals, restatements, and post-IPO firm performance); and (4) we examine the long-run economic consequences of earnings management in VC-backed IPOs. In particular, we expect that IPOs will exhibit more positive long-term stock performance in cases where VCs constrain opportunistic earnings management.

Overall, our results indicate that VC presence and incentives to monitor are associated with lower earnings management. Our findings are robust to controls for the endogenous choice of VC financing.[1] We consistently find that the VC effect holds even when controlling for the height of the Internet IPO bubble, IPO lock-up provisions, and VC partial cashing-out subsequent to the IPO. These findings are also robust to alternative proxies for earnings management. Finally, we provide some evidence that the lower earnings management associated with VC monitoring partially explains the superior post-IPO returns of VC-backed firms.

The paper proceeds as follows. Section II provides a summary of the institutional background, as well as the extant literatures that separately examine VCs and earnings management in IPOs. We present the formal hypotheses in Section III. The sample and research design are described in Section IV, and results are detailed in Section V. Conclusions and suggestions for future research are found in Section VI.

II. LITERATURE REVIEW

Overview of the Venture Capital Structure

A VC organization generally comprises various investment "funds." A fund represents capital raised by the VC from passive investors for a specific investment strategy and is governed by extensive contract details intended to protect the interests and tax reporting needs of the passive investors. The fund itself is most often structured as a limited liability partnership (LLP) with a limited life of 10–12 years, after which the LLP is completely dissolved and the original capital plus any contractual returns on the investments are repatriated to the limited partners (i.e., the passive investors). General partners retain what is left, move on to raising new "funds," create new LLPs, and manage their remaining portfolio of investments. A VC organization might have several funds under its management at any given time, and its long-run success is driven in part by its maintained ability to raise new investment funds from passive investors. This "fund-raising" success rests on the VC's reputation for returning profits to its passive investors that are higher than in the open market (Gompers and Lerner 2002, 242).

The Venture Capital Cycle

The VC fund generally follows three stages: (1) pre-screening of business plans and company selection; (2) nurturing and monitoring their investments; and (3) harvesting the investment—whether through an IPO, a merger, or a sale—and then exiting (see, Gompers and Lerner [2002] for a detailed discussion of the VC cycle).

Pre-Screening and Company Selection

The investment strategies of a VC organization can vary by fund, or the VC can limit itself to one strategy pursued across multiple funds. For example, VCs can specialize in one industry or be generalists across multiple types of firms (Sahlman 1990; Barry et al. 1990; Kaplan and Strömberg 2003). Once the VC has raised the cash and structured a fund, the organization actively pursues the agreed-upon investment goals of the fund. This process begins with extensive research, due diligence, and other pre-investment screening activities that are considered key to a VC's long-run investment success (Kaplan and Strömberg 2000). After the VC decides to invest in a particular company, fairly standard contract terms with the investee are observed across VC firms. These contracts include specification of the following: active ongoing monitoring and information requirements, often involving regular detailed reporting; the right to replace or hire new management; a certain number of seats on the board of directors that often increase with the amount of cash invested; and compensation, exit, and liquidation rights (Barry et al. 1990; Gompers 1995; Kaplan and Strömberg 2001; Gompers and Lerner 2002, Chapter 8).

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[1] Prior research and anecdotal industry evidence suggest that VC financing is an endogenous choice by both the entrepreneurs and the VCs (e.g., Lerner 1995; Gompers and Lerner 1998a; Hsu 2004; Lee and Wahal 2004). This choice is reflected in the nonrandom distribution and characteristics of VC-backed IPO firms versus non-VC-backed IPO firms. We use sophisticated matching methods to endogenize VC presence in an IPO and present both univariate and multivariate results. Our findings are robust across three different methods of one-to-one matching. A more detailed discussion appears in the Appendix.

Nurturing and Monitoring Investments

In general, VCs target firms where they can provide significant value-added services and expertise, not just cash infusions. This approach allows them to acquire equity at a significant discount (Hsu 2004). VCs can be involved in virtually all aspects of a firm's management, especially in the early years of a firm and/or as the firm moves closer to an IPO, merger, sale of the firm, etc. For instance, VCs help the portfolio management to hire key employees, introduce the portfolio companies to key suppliers and customers, and help to develop business strategies (for an overview, see, Sahlman 1990; Barry et al. 1990; Hellman and Puri 2002; Gompers and Lerner 2002, Chapter 8). Such close proximity to the firms' day-to-day operations, including financial reporting decisions, by a significant and often powerful investor can reduce the pre-IPO information asymmetry (Sahlman 1990; Black and Gilson 1998).

Harvesting and Exiting the Investment

The VC harvests the investment via one of the following choices: IPO, merger, sale, or liquidation of the firm. Selling the portfolio firm via a merger or acquisition deal is the most frequent form of harvesting by VCs (23.8 percent). Only slightly less common are IPOs (22.5 percent). Despite being less common than merger and acquisition (M&A) exits, IPOs on average provide the greatest returns to the VC funds of all possible exit strategies (Gompers and Lerner 2002, 150). Liquidation occurs infrequently and is usually reserved for underperforming firms.

Rarely do VCs complete a full exit at the IPO by selling their shares on the open market—i.e., by "cashing out." Barry et al. (1990) document that VCs sell approximately 5.2 percent of their pre-IPO shares and suggest that the IPO should not be viewed as an immediate exit strategy. Darrough and Rangan (2005) show that, on average, VCs sell only about 1 percent of the pre-IPO shares at the time of the IPO. Venture capitalists hold a vast majority of their original investment for at least the duration of the lock-up period required by IPO underwriters. Lock-up periods are an attempt to align the original shareholders' interests with the interests of the new IPO shareholders by preventing any sale or distribution of the stock by the original shareholders for a proscribed period—usually 180 days.

Gompers and Lerner (2002, 266) observe that, on average, VCs hold a majority of their shares for two years after the initial offering. VCs seldom cash out their remaining holdings even after the expiration of the lock-up period. Rather, they typically *distribute* the shares to the limited partners. They use the market value at the distribution date to determine the dollar value of their repatriation of the passive investors' initial investment and subsequent contracted percentage of the profits (Gompers and Lerner 1998b; Gompers and Lerner 2002, 209–210).[2] Also, VCs often retain a presence on the board of directors for many years after the lock-up expiration (Gompers and Lerner 2002, Chapter 13; Barry et al. 1990; Heesen 2003).

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[2] Gompers and Lerner (1998b) provide descriptive statistics to show that only 1 percent of the distributions occur in the three months immediately after going public. Note that data on these distributions made by VCs to their limited partners are not publicly available.

Venture Capitalist Reputation and IPOs

Venture capitalists are highly motivated to be associated only with successful and credible IPO companies due to additional scrutiny and public attention that companies receive when they first go public. This might motivate VCs to use their proximity to and influence over management to impact the portfolio firm's accounting and reporting decisions—in particular, the earnings management decisions (Heesen 2002). In addition, VCs often cite the profitable companies they have brought public in their marketing documents distributed to the passive investors (Gompers and Lerner 1997). Hence, successful IPO backing aids VCs in their future fundraising efforts and improves the ability of VCs to take firms public in the future.

Do Venture Capitalists' Interests Conflict with Those of Other Shareholders?

Although the VC literature provides evidence that VCs add value, some studies suggest that in certain cases the interests of VCs may conflict with those of the other stakeholders—i.e., the "grandstanding theory" where young or lower quality VCs prematurely take firms public (Gompers 1996; Lee and Wahal 2004). In an IPO, the pressures on management to have a successful offering are likely aligned with the VCs' pressures to provide high returns to their own investors (i.e., their limited partners). Because VCs must by design eventually exit the IPO to provide a return to their own limited-partner investors, they could indeed have short-run incentives that conflict with the new IPO shareholders or the other pre-IPO stakeholders. Cashing out their interests shortly before or after an over-hyped or overvalued IPO is another example of how the interests of VCs might conflict with those of the other stakeholders.

Recent empirical work in the accounting literature offers a mixed view of the role of VCs and accounting information. Cohen and Langberg (2005) find that VC presence, equity stake, and board seats are negatively related to earnings reactions and earnings response coefficients for their samples of firms. They conclude that VC backing is costly to public firms and that VCs' investment horizons lead them to use their ownership status in a myopic and self-serving manner. However, their study does not focus on actual management reporting behavior that might be affected by VC presence, or whether the market simply requires or expects less information from VC-backed firms. Conversely, Felix et al. (2005a, 2005b) provide evidence that VCs are perceived by investors to provide information and certification incremental to that of the external auditors as reflected in the first-day IPO underpricing setting. They also find that the presence of VCs attenuates investor demand for insurance from the external audit function.

Earnings Management in IPOs

Prior IPO earnings management research indicates that many IPO firms exhibit "unusually high" income-increasing abnormal accruals in both the pre-issue and post-issue time periods (Teoh et al. 1998b; Teoh et al. 1998; Aharony et al. 1993). More importantly, increasing levels of these accruals are associated with poorer long-run stock return performance (Teoh et al. 1998b). However, existing studies of earnings management have not yet documented potential constraints to IPO earnings management and whether these constraints might be associated with improved firm performance, thereby enhancing shareholder wealth.

III. HYPOTHESES DEVELOPMENT

Prior studies suggest that IPO firms exhibit upward earnings management at the time of the offering to temporarily inflate the stock offering price (Teoh et al. 1998b; Teoh et al. 1998; Dechow and Skinner 2000). Studies also demonstrate that VCs play a general monitoring role (e.g., Megginson and Weiss 1991). We extend prior studies to examine the monitoring role of VCs within a financial reporting context in an IPO. We do not predict *zero*, but rather *lower*, earnings management in VC-backed IPOs. Not all earnings management is "bad" (e.g., Subramanyam 1996; Shivakumar 2000). A monitor could conceivably use "good" earnings management to smooth income. Since our hypothesis is focused on VCs constraining "bad" earnings management, we structure our tests and additional analyses to ensure that we capture this type of earnings management.

More formally:

H1: We expect IPO-year abnormal accruals to be lower for VC-backed IPOs, relative to non-VC-backed IPOs.

IV. SAMPLE AND RESEARCH DESIGN

Sample Selection

We obtain our initial sample of domestic U.S. IPOs from the Securities Data Company (SDC) IPO and VentureXpert databases for the period 1983–2001. We use 2001 as the cutoff year because some of our tests require three years of data subsequent to the IPO year. For inclusion in the final sample, IPOs must have the available Compustat financial data both in the year of and the year prior to the offering to calculate discretionary accruals, as well as stock return data during the three-year period beyond four months after the offering. These selection criteria yield an initial sample size of 3,090 IPO firms. We exclude from our sample firms that engaged in merger and acquisition activity in the pre-IPO period, since Hribar and Collins (2002) suggest that such activities can introduce measurement error to our abnormal accruals measures (167 observations). Furthermore, the top and bottom 5 percent of outlier values calculated for discretionary accruals (discussed in detail later in this section) are dropped from the sample, yielding a final sample of 2,630.

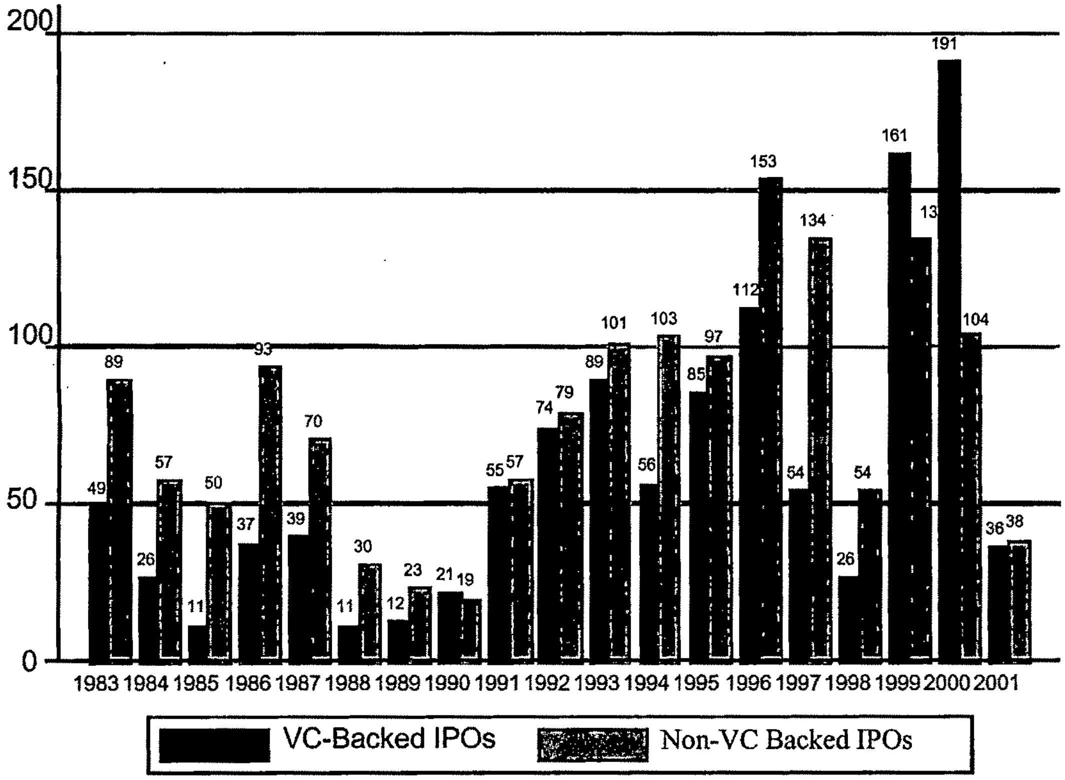
Research Design

Our research design controls for endogeneity in the decision of VCs to back an IPO firm in both our univariate and multivariate tests via a matching framework and an endogenous switching regression, respectively. Extant literature finds that venture financing represents an endogenous choice made by entrepreneurs and VCs (Lee and Wahal 2004). The outcome of this choice is observed in the nonrandom distribution and characteristics of VC-backed IPOs. Lerner (1995) and Lee and Wahal (2004) find that VC-backed IPOs show clustering across both industry and geographical dimensions. Lee and Wahal (2004) also find that VCs tend to invest in younger, smaller, less-profitable firms with lower book values of equity, lower revenues, and lower total assets.

Figure 1 provides a comparison of the venture-backed IPOs against the total IPOs by year over the period of interest. Approximately 43 percent of IPOs are backed by VCs over the period 1983–2001, suggesting that VC presence plays a material role in facilitating the process of going public. This is consistent with the documentation of prior studies. However, Figure 1 also shows that there is significant variation in VC presence in IPOs across years, ranging from 18 percent of all IPOs in 1985 to 65 percent in 2000 at the height of the dot-com IPO bubble.

Table 1 provides the descriptive statistics for the 2,630 IPO firms in our sample. Panel A presents the distribution of IPO firms by industry, using two-digit SIC codes, and further

FIGURE 1
Number of Venture Capital-Backed (VC) Initial Public Offerings and Non-Venture Capital-Backed (Non-VC) Initial Public Offerings during the Sample Period 1983–2001
(n = 2,630)



partitions each industry into subsamples of VC-backed and non-VC-backed IPOs.[3] The industry distribution of IPO firms is similar to that of Teoh et al. (1998b). Table 1, Panel A, also shows that approximately 63 percent of VC-backed IPOs tend to cluster in four industries—chemical products, scientific instruments, computer hardware and software, and health. Our data also shows geographic concentrations. In Table 1, Panel B, we show VC-backed IPOs as a percentage of all IPOs for the state in which the firm is headquartered. For brevity, we show this distribution only for states that represent at least 2 percent of all IPOs during the period 1983–1998 (i.e., at least 61 IPOs). More than half of the IPOs in California and Massachusetts are backed by VCs. These two states alone represent approximately 48 percent of all VC-backed IPOs. Lerner (1995) suggests that VCs tend to invest in companies close to their own headquarters; since a significant number of VCs tend to be based in California and Massachusetts, this geographic clustering is sensible.

Table 1, Panel C, presents the firm and offer characteristics partitioned by the type of IPO (VC-backed or non-VC-backed IPO). Venture capitalists tend to invest in particular

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[3] The grouping of SIC codes into industry name categories is consistent with Teoh et al. (1998) to maintain comparability.

TABLE 1
Sample Descriptives for 2630 IPOs from 1983–2001

Panel A: SIC Distribution

Industry	Two-Digit SIC Code	No. Firms	Percent of Total IPOs	No. VC-Backed	Percent of VC-Backed IPOs in Industry
Oil and Gas	13, 29	23	0.87	6	26.09
Food Products	20	48	1.83	12	25.00
Paper and Paper Products	24–27	63	2.40	19	30.16
Chemical Products	28	159	6.05	101	63.52
Manufacturing	30–34	104	3.95	20	19.23
Computer Hardware and Software	35, 73	811	30.84	454	55.98
Electronic Equipment	36	269	10.23	128	47.58
Transportation	37, 39, 40, 42, 44, 45	116	4.41	16	13.79
Scientific Instruments	38	225	8.56	127	56.44
Communications	48	124	4.71	59	47.58
Electric and Gas Services	49	33	1.25	11	33.33
Durable Goods	50	73	2.78	15	20.55
Retail	53, 54, 56, 57, 59	117	4.45	34	29.06
Eating and Drinking Establishments	58	38	1.44	11	28.95
Financial Services	61, 62, 64, 65	35	1.33	7	20.00
Entertainment Services	70, 78, 79	51	1.94	9	17.65
Health	80	80	3.04	41	51.25
All Others	10, 14, 15, 16, 17, 22, 23, 47, 51, 52, 55, 63, 67, 72, 75, 82, 83, 87, 99	261	9.92	74	28.35
Total		2630	100.00	1144	43.50

Panel B: Geographic Distribution

State	No. VC-Backed IPOs (% of total for state)	No. Non-VC-Backed IPOs (% of total for state)	Total No. IPOs (% of total of all IPOs)
California	431 (62.46)	259 (37.54)	690 (26.27)
Colorado	19 (32.20)	40 (67.80)	59 (2.25)
Florida	27 (26.47)	75 (73.53)	102 (3.88)
Georgia	34 (46.58)	39 (53.42)	73 (2.78)
Illinois	28 (34.57)	53 (65.43)	81 (3.08)
Massachusetts	118 (64.48)	65 (35.52)	183 (6.97)
Minnesota	31 (43.06)	41 (56.94)	72 (2.74)
New Jersey	35 (33.02)	71 (66.98)	106 (4.04)
New York	45 (24.19)	141 (75.81)	186 (7.08)
Pennsylvania	37 (44.05)	47 (55.95)	84 (3.20)
Texas	67 (38.73)	106 (61.27)	173 (6.59)

(*continued on next page*)

TABLE 1 (Continued)

Panel C: Firm Characteristics (per share basis unless otherwise stated): Means and (Medians)

	VC-Backed IPOs Difference in Means and (Medians) (n = 1486)	Non-VC-Backed IPOs Difference in Means and (Medians) (n = 1144)	t-statistic (z-statistic)[a]
Sales	5.620	9.649	−8.544***
	(2.555)	(5.501)	(−12.387)***
Total Assets	6.652	8.723	−3.229***
	(4.504)	(5.395)	(−3.999)***
Net Income	−0.277	−0.035	−0.870
	(−0.127)	(0.303)	(−12.387)
Book Value of Equity	3.496	3.726	−1.672*
	(3.272)	(3.133)	(1.572)
Earnings (Operating)	46.94	70.44	−12.537***
(% positive indicator variable)	(0.000)	(1.000)	(−12.181)***

Panel D: Use of IPO Proceeds[b]

	Non-VC	% of Non-VCs	VC	% of VCs	Difference in % (Non-VC % − VC %)	Total	% of Total
Current and future acquisitions	164	7.88	154	9.00	−1.12	318	8.39
Capital Expenditure	69	3.32	72	4.21	−0.89	141	3.72
General Purpose	765	36.78	656	38.34	−1.56	1421	37.48
Investments	26	1.25	29	1.69	−0.44	55	1.45
Operations	45	2.16	67	3.92	−1.75	112	2.95
Other	13	0.63	3	0.18	0.45	16	0.42
Paying down on borrowings	127	6.11	60	3.51	2.60	187	4.93
Product development/R&D	39	1.88	75	4.38	−2.51	114	3.01
Refinancing/Retire debt	664	31.92	362	21.16	10.77	1026	27.06
Stock repurchases	54	2.60	39	2.28	0.32	93	2.45
Working capital	114	5.48	194	11.34	−5.86	308	8.12
Total	2080	100.00[c]	1711	100.00[c]		3791	100.00[c]

Panel E: Control Variables

	VC-Backed IPO Difference in Means (Medians) (n = 1486)	Non-VC-Backed IPO Difference in Means (Medians) (n = 1144)	t-statistic (z-statistic)[a]
AUDQUAL	0.969	0.790	13.876***
	(1.000)	(1.000)	(13.397)***
UWRITER	7.305	6.161	9.508***
	(8.100)	(7.100)	(12.532)***
GROWTH	2.365	1.271	1.794*
	(0.541)	(0.305)	(9.714)***

(*continued on next page*)

TABLE 1 (Continued)

Panel E: Control Variables (Continued)

	VC-Backed IPO Difference in Means (Medians) (n = 1486)	Non-VC-Backed IPO Difference in Means (Medians) (n = 1144)	t-statistic (z-statistic)[a]
LEV	0.112	0.126	−1.686*
	(0.015)	(0.0308)	(−3.688)***
LNTA	4.034	3.715	5.457***
	(4.026)	(3.651)	(7.363)***
LMVE	4.756	4.105	10.107***
	(4.706)	(4.086)	(9.980)***
PERFORM	0.781	0.778	0.179
	(1.000)	(1.000)	(0.191)

***, * Denote two-tailed significance at the 1 percent and 10 percent levels, respectively.

The sample consists of domestic IPO firms going public in the period 1983–2001. The sample firms must have the necessary Compustat data to calculate the abnormal current accruals using the cross-sectional modified Jones model as applied by Teoh et al. (1998) and the performance-matched accruals as applied by Kothari et al. (2005).

Panel A reports the distribution of IPO firms by two-digit SIC code. The industry name grouping is based on Teoh et al. (1998).

Panel B tabulates the geographic distribution of the IPO firms.

Panel C provides mean (median) firm and offer characteristics.

Panel D tabulates the use of IPO proceeds by the firm.

Panel E provides mean (median) information on the control variables used in Model 1.

[a] Mann-Whitney tests are performed.

[b] The use of IPO proceeds is identified from the SDC Initial Public Offerings database. Each IPO firm may have multiple uses of the IPO proceeds; hence, the total number exceeds 2630 IPOs.

[c] Percentages do not sum to 100 percent due to rounding.

$AUDQUAL_i$ = an indicator variable that equals 1 if the auditor is a Big 6 auditor for firm *i*, and 0 otherwise; this variable is defined to take into account the impact of the various mergers among the leading audit firms in the 1980s and 1990s;
$UNDERWRI_i$ = firm *i*'s underwriter ranking using the Caster-Manaster reputation rankings for IPO underwriters;
$GROWTH_i$ = change in sales between the year of and year prior to the offering for firm *i*, scaled by the sales in the year prior to the offering;
LEV_i = long-term debt scaled by total assets for firm *i*;
$LNTA_i$ = natural logarithm of total assets for firm *i*; and .
$PERFORM_i$ = an indicator variable that equals 1 if the funds flow from operation (Compustat data 110) or cash flow from operations (Compustat data 308) for firm *i* is greater than the two-digit SIC code industry average, and 0 otherwise.

types of firms: those that are relatively small (as measured by total assets per share), those that feature lower sales, and those that generate losses (as scaled by number of shares) (t = −3.229, t = −8.544, and t = −12.537, respectively). Overall, these firm characteristics are consistent with prior literature and suggest that the non-randomness of the data can be used to instrument the VC financing choice (Lee and Wahal 2004). We report our main findings with and without the endogeneity controls to provide evidence on the importance of this control for subsequent research regarding VC effects on public firms' accounting decisions. The Appendix provides a detailed discussion of our methods for controlling for endogenous choice.

Table 1, Panel D, presents the details on the use of IPO proceeds. Overall, we observe that non-VC-backed IPO firms tend to use the proceeds from the IPO to refinance or retire debt. Conversely, VC-backed IPO firms tend to use the IPO proceeds for working capital and product/research development.

Venture Capitalists and Earnings Management

To examine the association between VCs and earnings management, we regress abnormal current accruals on our variable of interest—an indicator variable that captures the presence or absence of a VC in the IPO (*VCIPO*) and other control variables (see H1). If VCs are associated with lower abnormal accruals around the IPO, then we would expect *VCIPO* to be significantly negative.

Research Design Issues

In our research design, we acknowledge and attempt to adjust for the inherent difficulties associated with measuring earnings management via discretionary accruals models (see Dechow et al. 1995; Kothari et al. 2005). A key observation in Dechow et al. (1995) is that the models are not well specified for differentiating between "abnormal" accruals that are consistent with growing firms and those that are consistent with earnings management.

We address this issue in several ways in our research design. First, we introduce a control proxy for growth, which we describe in the subsection below. Second, we rerun our tests and use the performance-matching approach of Kothari et al. (2005) to compute abnormal discretionary accruals. This method attempts to rigorously match firms in order to minimize the positive correlation between extreme firm performance and the abnormal-accruals measures. Finally, by partitioning the sample on VC-backed IPOs, we add yet another control for both growth and the life cycle of the IPO firm, since VCs generally invest in high-growth entrepreneurial firms (Kaplan and Strömberg 2004).[4] All of these factors generally bias against our finding the predicted negative association between our variable of interest (VC presence) and abnormal accruals.

Regression Model

Our regression model (Model 1) is as follows:

$$DCA_i = \alpha_0 + \beta_1 VCIPO_i + \beta_2 AUDQUAL_i + \beta_3 UNDERWRI_i + \beta_4 GROWTH_i + \beta_5 LEV_i + \beta_6 LMVE_i + \beta_7 PERFORM_i + \varepsilon_i \quad (1)$$

where:

DCA_i = abnormal accruals calculated using either (1) the cross-sectional adaptation of the modified Jones (1991) model, or (2) the performance-matched abnormal-accruals model (Kothari et al. 2005) to obtain the discretionary and nondiscretionary accruals for firm *i* for a given year;

$VCIPO_i$ = an indicator variable that equals 1 if the IPO is backed by a VC for firm *i*, and 0 otherwise;

$AUDQUAL_i$ = an indicator variable that equals 1 if the auditor is a Big 6 auditor for firm *i*, and 0 otherwise; we account for the impact of the various mergers among the leading audit firms in the 1980s and 1990s;

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[4] This partition is consistent with the spirit of the concerns raised in prior research by McNichols and Wilson (1988), Dechow et al. (1995), and McNichols (2000). They note that in many studies of earnings management researchers partition the sample such that the variable of interest and the accruals measures are correlated in the same direction—i.e., the partition choice biases in favor of finding predicted results. Our research design directly addresses that concern and biases against finding our hypothesized relationships.

$UNDERWRI_i$ = firm i's underwriter ranking using the Carter et al. (1998) reputation rankings for IPO underwriters (these rankings vary from one to nine, and perceived reputation increases with the ordinal ranking);[5]

$GROWTH_i$ = change in sales between the year of and year prior to the offering for firm i, scaled by the sales in the year prior to the offering;

LEV_i = long-term debt scaled by total assets for firm i;

$LMVE_i$ = natural logarithm of market value of firm i at fiscal year-end; and

$PERFORM_i$ = an indicator variable that equals 1 if the funds flow from operations (Compustat data 110) or cash flow from operations (Compustat data 308) for firm i is greater than the industry average for the two-digit SIC code.

Calculating Abnormal Accruals

We adopt the approach used in prior IPO-related earnings management research (e.g., see Teoh et al. 1998a, 1998b) using the cross-sectional adaptation of the modified Jones (1991) model to obtain the discretionary and nondiscretionary accruals for a given year. Similar to Teoh et al. (1998b), we focus on discretionary current accruals (*DCA*) as the key variable that captures earnings management. Teoh et al. (1998) find evidence that current abnormal accruals have the greatest consistent explanatory power among all the proxies for earnings management examined in an IPO setting.

Current accruals (*CA*) are defined as the difference between the change in noncash current assets and the change in operating current liabilities. In terms of Compustat data items, this is the change in [accounts receivables (#2) + inventory (#3) + other current assets (#68)] − change in [accounts payable (#70) + tax payable (#71) + other current liabilities (#72)]. The nondiscretionary accruals for an IPO firm in a given year is estimated from a cross-sectional regression in that year of *CA* scaled by lagged total assets on the change in sales scaled by lagged total assets:

$$CA_{i,t}/TA_{i,t-1} = \alpha_0\left(\frac{1}{TA_{i,t-1}}\right) + \alpha_i\left(\frac{\Delta Sales_{i,t}}{TA_{i,t-1}}\right) + \varepsilon_{i,t}. \tag{2}$$

We estimate the Jones regression cross-sectionally for a sample of all available firms in Compustat in the same two-digit SIC code as the IPO firm, excluding the IPO firm from this sample. The cross-sectional regressions are re-estimated for each year in the period 1983–2001. At least ten observations must be available in order to perform the regression. Hence, the nondiscretionary current accruals are calculated as:

$$NCA_{i,t} = \hat{\alpha}_0\left(\frac{1}{TA_{i,t-1}}\right) + \hat{\alpha}_1\left(\frac{\Delta Sales_{i,t} - \Delta Trade\ Receivables_{i,t}}{TA_{i,t-1}}\right). \tag{3}$$

The discretionary current accruals are calculated as the difference between *CA* and *NCA*.

We also employ the performance-matched discretionary accruals approach that Kothari et al. (2005) suggest. This approach is especially relevant in the VC-backed IPO setting, because performance might be a key driver of both a firm's decision to go public and of a VC's decision to provide financial backing to an IPO firm. By incorporating this performance-matched approach in our research design when estimating discretionary accruals, we attempt to control for systematic differences in performance (and growth) between VC-backed and non-VC-backed IPOs.

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[5] We also use the updated underwriter rankings for more recent years that have been generously provided by Jay Ritter on his website: http://bear.cba.ufl.edu/ritter/.

Similar to Kothari et al. (2005), we estimate the performance-matched modified Jones discretionary accrual as the difference between the cross-sectional modified Jones model discretionary accrual and the corresponding discretionary accrual for a performance-matched firm. We use the current-year return on assets—ROA, measured by the ratio of net income and lagged total assets—as the performance measure. We match each IPO firm with another non-IPO Compustat firm in the same two-digit industry and year with the closest ROA in the current year. In untabulated descriptive results, we find that VC-backed and non-VC-backed firms exhibit different performance levels. Return on assets, sales, and cash flow from operations (scaled by total assets) for our sample of VC-backed IPO firms are all significantly lower compared to our sample of non-VC-backed IPO firms ($t = -4.249$, $t = -4.461$, and $t = -2.702$, respectively).

Control Variables

External Monitors

With respect to external monitors, Aharony et al. (1993) test whether the joint quality of the auditor and underwriter constrains the extent of earnings management. With the reputations of the auditor and underwriter at stake, and considering that they could incur considerable litigation-based loss if they fail to reveal or prevent material misstatements, higher quality auditors and underwriters are likely motivated to detect "bad"earnings management. This leads to the expectation (not supported by Aharony et al. [1993]) that the extent of earnings management is inversely related to the joint quality of the auditor and underwriter. To ensure that we fully examine the potential impact of VCs incremental to these other key players or monitors, we examine separately the roles of auditor (*AUDQUAL*) and underwriter (*UNDERWRI*) quality.

Financial Variables

Consistent with prior research, we also control for size (*LMVE*), growth (*GROWTH*), and financial structure (*LEV*) of the firm. If growth or firm size affects abnormal current accruals, we expect significantly positive coefficients. We control for financial structure because a highly leveraged firm faces high financial or bankruptcy risk that could jeopardize the success of the IPO and encourage earnings management to avoid covenant default around the IPO. Alternatively, highly levered firms face greater monitoring from debt holders. As such, we do not have a specific expectation for the coefficient on *LEV*.

We then extend prior models by introducing a variable intended to capture the cash flow performance of the IPO firm relative to the rest of its peers in the industry. It is possible that IPO firms backed by VCs are, on average, "better performers" than IPO firms not backed by VCs and therefore do not need to manage earnings. These better performers (in terms of cash flows) would exhibit lower abnormal accruals due to performance or to VC pre-investment screening, not due to VC monitoring. We add a non-accruals-based performance measure (*PERFORM*), as an indicator variable, set equal to 1 if cash flows from operations is greater than the industry average, and 0 otherwise. We also include industry and year dummies in the regressions (results not tabulated).

Table 1, Panel E, presents the descriptive statistics for our control variables. VC-backed IPOs tend to have higher quality auditors and underwriters, higher growth, and lower leverage.

V. EMPIRICAL RESULTS

Earnings Management in the Presence of VCs

Univariate Results

Table 2, Panel A, presents the summary statistics for abnormal accruals for the overall sample without adjusting for endogeneity concerns. Depending on which accruals model we use, between 63 percent and 65 percent of all the IPO firms in the sample exhibit positive abnormal accruals in the year of the offering. This is consistent with the IPO sample in Teoh et al. (1998b). Testing whether the mean abnormal accrual is different from zero yields significant t-statistics for the two accruals models used. This is initial evidence of systematic upward earnings management activity for these IPO firms. These results are consistent with the conjecture that firms manage their earnings upward when first accessing the public equity markets (Teoh et al. 1998b; Teoh et al. 1998; Dechow and Skinner 2000).

Table 2, Panel B, shows that VC-backed IPO firms generally exhibit significantly lower (mean and median) abnormal accruals compared with non-VC-backed IPO firms. The univariate results provide initial evidence (at the 1 percent level of significance) that VCs mitigate upward earnings management. We also note that Table 2, Panel B, shows that both VC-backed and non-VC-backed IPOs exhibit earnings management behavior. These results are consistent with prior studies that find that earnings management is expected on average in IPOs (Teoh et al. 1998). Again, we do not make the strong conjecture that VC presence *eliminates* earnings management entirely, but instead hypothesize that VC presence likely *mitigates* "bad" earnings management behavior.

Table 2, Panel C, presents the endogeneity-adjusted univariate results for the difference in abnormal accruals between VC-backed and non-VC-backed IPOs. Each VC-backed IPO is matched with one or more non-VC-backed IPO(s) using the highest propensity score, Gaussian kernel weighting, and regression-adjusted local-linear methods (see the Appendix). After we adjust for the endogenous choice for VC financing, VC-backed IPOs continue to exhibit significantly lower abnormal accruals across all of the three matching approaches we adopt.

Overall, we find that the presence of a VC is significantly associated with decreased earnings management in IPO firms. These results are robust across various model specifications and alternative definitions of earnings management, as measured by the use of abnormal accruals and adjustments for endogeneity using a variety of matching methods.

Multivariate Results

Main Hypothesis

The structure of our multivariate tests and sensitivity analyses is as follows: (1) we estimate the OLS regression for Model 1, which examines whether VC presence and incentives to monitor are associated with lower abnormal accruals, per H1; (2) we perform additional tests to ensure that our main results are actually driven by VCs and not by competing explanations (e.g., the Internet IPO bubble, lock-ups, the endogenous choice of VC financing, or subsequent VC exits); (3) we confirm that our findings are robust across various earnings management measures (e.g., cross-sectional-Jones and performance-matched abnormal accruals, restatements, or post-IPO firm performance); and (4) we document the differing effects of earnings management on the long-run returns in VC-backed versus non-VC-backed IPOs.

Table 3 presents the regression results for H1—both adjusted and unadjusted for the endogenous choice of VC financing. The relevant test variable is *VCIPO*; we expect the coefficient on this variable to be significantly negative if the VC is acting as an additional monitor. First, we turn our attention to the unadjusted (for endogenous choice) results. We

TABLE 2

Univariate Tests of Abnormal Accruals—With and Without Endogeneity Adjustments

Panel A: Abnormal Accruals (% of Total Assets) Total Sample (n = 2630)

	Mean	Median[a]	25th Percentile	75th Percentile	Percentage Positive
Cross-Sectional Modified Jones Abnormal Accruals	5.132***	4.131***	−5.678	18.857	62.62
Performance-Matched Cross-Sectional Modified Jones Abnormal Accruals	4.532***	3.194***	−9.351	21.596	64.71

Panel B: Abnormal Accruals (% of Total Assets) by VC- and Non-VC-Backed IPOs (not adjusted for endogeneity)

	VC-Backed IPO Means (Medians)[a] (n = 1486)	Non-VC-Backed IPO Means (Medians)[a] (n = 1144)	Difference in Mean (Median) t-statistic (z-statistic)[b]
Cross-Sectional Modified Jones Abnormal Accruals[c]	3.619*** (3.312)***	6.298*** (4.777)***	−1.886* (−2.270)**
Performance-Matched Abnormal Accruals[c]	2.660** (2.223)***	6.000*** (4.365)***	−2.080** (−2.682)***

Panel C: Abnormal Accruals (% of Total Assets) by VC- and Non-VC-Backed IPOs (adjusted for endogeneity[d])

	Propensity Score Difference in Means between VC-Backed and Non-VC-Backed IPOs (t-statistic)	Gaussian Kernel Difference in Means between VC-Backed and Non-VC-Backed IPOs (t-statistic)	Local-Linear Regression-Adjusted Difference in Means between VC-Backed and Non-VC-Backed IPOs (t-statistic)
Cross-Sectional Modified Jones Abnormal Accruals	−0.033 (−2.554)**	−0.023 (−2.023)**	−0.021 (−1.877)*

***, **, * Denote two-tailed significance at the 1 percent, 5 percent, and 10 percent levels, respectively.

t-statistics for differences in means appear in parentheses.

Two-digit SIC code dummies, year dummies, headquarter state dummies, book value per share, sales per share, and total assets per share are used as instrumental variables.

[a] Sign test that the true proportion of positive (negative) signs is one-half.

[b] Mann-Whitney tests are performed.

[c] The abnormal current accruals are calculated using the (1) cross-sectional modified Jones (1991) model to obtain the discretionary and nondiscretionary accruals for firm *i* for a given year; or (2) performance-matched abnormal accruals (cross-sectional modified Jones-model abnormal accrual—matched firm's cross-sectional modified Jones-models abnormal accrual) per Kothari et al. (2005).

[d] Endogeneity-adjusted abnormal accruals differences between VC-backed and non-VC-backed IPOs are presented. Each VC-backed IPO is matched with one or more non-VC-backed IPOs using the propensity (nearest propensity score), Gaussian kernel, and regression-adjusted local-linear matching approaches.

note significant results across the two proxies for earnings management (modified Jones and performance-matched abnormal accruals) that are consistent with the VC monitoring hypothesis in this setting. After controlling for other variables, the coefficient on *VCIPO* is

TABLE 3
Venture Capitalist Presence and Earnings Management Regression Results

Variable	Expected Sign	Unadjusted for Endogenous Choice of VC Financing: Modified Jones DCA	Unadjusted for Endogenous Choice of VC Financing: Performance-Matched DCA	Adjusted for Endogenous Choice of VC Financing: Modified Jones DCA	Adjusted for Endogenous Choice of VC Financing: Performance-Matched DCA
Intercept	+	0.160	0.193	0.154	0.186
		(9.00)***	(8.09)***	(8.21)***	(7.36)***
VCIPO	−	−0.023	−0.031	−0.014	−0.017
		(−2.43)**	(−2.48)**	(−2.19)**	(−2.02)**
AUDQUAL	−	−0.000	0.011	−0.001	0.009
		(−0.03)	(0.55)	(−0.09)	(0.44)
UNDERWRI	−	−0.008	−0.009	−0.008	−0.009
		(−4.08)***	(−3.75)***	(−3.97)***	(−3.71)***
GROWTH	+	0.003	0.000	0.000	−0.000
		(0.43)	(0.04)	(0.44)	(−0.04)
LEV	No sign	−0.067	−0.069	−0.065	−0.070
		(−2.76)***	(−2.17)**	(−2.61)***	(−2.11)***
LMVE	+	−0.005	−0.004	−0.006	−0.004
		(−1.49)	(−0.93)	(−1.63)	(−0.91)
PERFORM	No sign	−0.008	−0.050	−0.008	−0.052
		(−0.72)	(−3.33)***	(−0.64)	(−3.35)***
Adj. R^2		2.4%	2.6%	2.3%	2.5%
n		2630	2630	2630	2630

***, ** Denote two-tailed significance at the 1 percent and 5 percent, respectively.
White (1980)-consistent t-statistics are in parentheses.
The sample consists of 2,630 domestic IPO firms going public in the period 1983–2001. The sample firms must have the necessary Compustat data to calculate the abnormal current accruals using the modified Jones model as applied by Teoh et al. (1998b). The following regression is conducted:

$$DCA_i = \alpha_0 + \beta_1 VCIPO_i + \beta_2 AUDQUAL_i + \beta_3 UNDERWRI_i + \beta_4 GROWTH_i + \beta_5 LEV_i + \beta_6 LMVE_i + \beta_7 PERFORM_i + \varepsilon_i$$

where:

DCA_i = abnormal current accruals calculated using (1) the cross-sectional modified Jones (1991) model to obtain the discretionary and nondiscretionary accruals for firm *i* for a given year; and (2) cross-sectional modified Jones (1991) models performance-matched abnormal accrual (the difference between the modified Jones abnormal accrual and that for a matched firm);
$VCIPO_i$ = an indicator variable that equals 1 if the IPO is backed by a VC for firm *i*, and 0 otherwise;
$AUDQUAL_i$ = an indicator variable that equals 1 if the auditor is a Big 6 auditor for firm *i*, and 0 otherwise; this variable is defined to take into account the impact of the various mergers among the leading audit firms in the 1980s and 1990s;
$UNDERWRI_i$ = firm *i*'s underwriter ranking using the Caster-Manaster reputation rankings for IPO underwriters;
$GROWTH_i$ = change in sales between the year of and year prior to the offering for firm *i*, scaled by the sales in the year prior to the offering;
LEV_i = long-term debt scaled by total assets for firm *i*;
$LMVE_i$ = natural logarithm of total assets for firm *i*; and
$PERFORM_i$ = an indicator variable that equals 1 if the funds flow from operation (Compustat data 110) or cash flow from operations (Compustat data 308) for firm *i* is greater than the two-digit SIC code industry average. Industry and year dummies are included in the regressions but results are not tabulated. We adjust for the endogenous choice for VC financing by using two-digit SIC code dummies, headquarter state dummies, year dummies, book value of equity per share, total assets per share, and sales per share as instrumental variables in the first-stage regression. Based on the parameter estimates in the first-stage probit regression, we estimate the probability of VC financing and use this probability measure in the second-stage regression.

significant and negative, as predicted (for VCs that monitor), across both measures of abnormal accruals.

The modified Jones model coefficient is −0.023 (t = −2.43), and the performance-matched accruals coefficient is −0.031 (t = −2.48). Our findings suggest that VC-backed IPO firms exhibit abnormal current accruals of 2.3–3.1 percent of total assets lower than non-VC-backed IPO firms. These multivariate results provide additional evidence consistent with the notion that VCs play an effective and measurable role in constraining the extent of earnings management among IPO firms. However, they do not rule out the possibility that the significance is actually driven by the endogeneity of VC-backing.

Table 3 also presents the regression results after adjusting for the endogenous choice for VC financing. Following Lee and Wahal (2004), we employ a two-stage procedure described in Section IV and the Appendix, where the second-stage regression (Model 1) uses the estimates from the first stage to provide consistent estimates of the parameters. We replace the dummy variable for VC presence with the estimated probability of a firm being backed by a VC based on the first-stage regression in our main regression.[6]

Results from the unadjusted regressions are generally preserved after implementing the endogenous choice adjustments. The coefficient on *VCIPO* is significantly negative, both when we use the modified Jones abnormal accruals as the dependent variable in the regression (t = −2.19) and when we use the performance-matched abnormal accruals (t = −2.02). Conditioning on the variables used to model the receipt of venture financing, firms with the backing of VCs exhibit abnormal accruals that are significantly lower than their non-VC-backed counterparts. These findings confirm that the endogeneity of this choice is not driving the overall support for H1.

While not the focus of our study, the coefficient on the external monitor control variable in both regressions, underwriter quality (*UNDERWRI*), are significantly negative and consistent with theory, but not with Aharony et al. (1993). The auditor quality (*AUDQUAL*) indicator is insignificant for both regressions. The insignificant coefficient on auditor quality, while contrary to theory, is consistent with prior research.[7]

Our cash-related control for firm performance (*PERFORM*) is negative for the regression with performance-matched accruals as the dependent variable, suggesting that better cash performers are associated with lower earnings management. Firm size (*LMVE*) is not significant. Finally, the control for the growth of the firm (*GROWTH*) is not significant across either specification of the model.

The adjusted-R^2 for the performance-matched accruals regression is 2.4 percent and 2.6 percent for the modified Jones accruals and performance-matched accruals regressions, respectively. We also find similar results for *VCIPO* after adjusting for the endogenous choice for VC financing. The results for the control variables are similar to the regressions where we do not adjust for the endogenous choice for VC financing. The adjusted-R^2s are 2.3 and 2.5 percent for both regressions, when we incorporate the adjustment for endogenous choice. In summary, after adjusting for the endogenous choice for VC financing and providing extensive controls for growth, our results continue to support the suggestion that VCs appear to mitigate earnings management in IPOs.

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[6] We also employ Heckman's two-step consistent estimator. The results are very similar. The selectivity correction λ has coefficients that are significant at p-values of 0.000 for both the modified Jones and performance-matched accruals regressions, suggesting that selectivity correction is important in the model.

[7] When we run the regression without our VC variables, both *AUDQUAL* and *UNDERWRI* are insignificant. We also rerun our main tests to include the interaction of *VCIPO* × *AUQUAL* and *VCIPO* × *UNDERWRI*; our results show that the interaction terms are not significant.

Competing Explanations and Further Analyses

Table 4 presents an overview of results from tests of the following potential competing explanations and further analyses: (1) the incentives of VC to monitor as proxied by investment in the firm; (2) the timing of VC exits from the IPO company; (3) the influence of the equity market bubble on VC monitoring; (4) the instance when VCs do not mitigate earnings management; and (5) the role of lock-up provisions that inhibit pre-IPO investors' ability to sell their holdings. We also discuss additional, untabulated findings regarding within-sample variation of earnings management and VC traits such as age of the VC firm, IPO experience, board membership, etc.

Is it Really VC Presence and Monitoring?

VC Monitoring

If our main results are consistent with VCs acting as monitors, then we should expect to observe the amount of abnormal accruals to be decreasing in proxies for VCs' incentives and ability to monitor. Consistent with existing theories of control, VCs are more likely to intervene to provide value-added and monitoring services as the VC's equity stake increases (Kaplan and Strömberg 2004). As the VC's stake in the IPO firm increases—as a proportion of the total known amount invested in the firm pre-IPO—we expect the VC to be more vigilant in constraining earnings management behavior (Berle and Means 1932; Jensen and Meckling 1976). Baker and Gompers (2003) show that boards of firms backed by VCs have fewer insiders as directors and suggest that VCs create boards with greater independence and oversight. Consistent with Baker and Gompers (2003), we predict that the VC provides additional monitoring as it increases its representation on the board of the investee firm.

In contrast, Gompers' (1996) "grandstanding" theory of VC firm behavior in IPOs suggests that younger and less experienced VC firms are more likely to prematurely bring their portfolio firms public in order to establish a successful track record with their limited partners (investors). Consistent with this theory, we expect younger and less experienced VC firms to be less likely to constrain earnings management compared to older and more established VC firms. This prediction is driven by the notion that "grandstanding" VC firms are most interested in simply taking the company public, rather than in performing the traditional monitoring and advising roles. Similarly, as VCs gain more experience in bringing their portfolio firms public, we expect this VC learning curve to be reflected in the decline of earnings management behavior of the portfolio firms that they take public over time.

Model 1 (removing *VCIPO*) is estimated for the sample of VC-backed IPOs only. We replace the dummy variable *VCIPO* in Model 1 with four additional VC-specific variables: *VCTOT*, *VCBOARD*, *VCEXP*, and *VCAGE*. The variables are measured as follows:

$VCBOARD_i$ = the number of individuals from the VC firm on the board of directors of firm i;
$VCTOT_i$ = the ratio of the dollar amount invested by the VC in firm i and the total known dollar amount invested in firm i prior to the IPO;
$VCAGE_i$ = the age of the VC firm for firm i; and
$VCEXP_i$ = the number of IPOs for the VC firm prior to the IPO for firm i.

Our untabulated results indicate that *VCTOT* is negatively associated with abnormal accruals, consistent with VCs acting as monitoring agents, whose incentives to monitor increase with their at-risk investment. These results hold for both the modified Jones and

TABLE 4
Venture Capitalist Presence and Earnings Management Regression Results—Sensitivity Analyses

Variable	**VC × Investment (1)**	**VC × Exit (2)**	**VC × Bubble (3)**	**VC × Less Aggressive (4)**	**VC × Lock-Up (5)**
Intercept	0.124	0.129	0.172	0.250	0.128
	(4.46)***	(4.38)	(10.36)	(11.52)***	(4.54)***
VCIPO	−0.016	−0.034	−0.030	0.004	−0.046
	(−1.03)	(−2.12)**	(−2.92)***	(0.25)	(−1.93)*
AUDQUAL	−0.017	−0.018	0.002	0.015	−0.018
	(−0.73)	(−0.73)	(0.11)	(0.81)	(−0.77)
UNDERWRI	−0.002	−0.002	−0.002	0.001	−0.002
	(−0.80)	(−0.71)	(−1.09)	(0.24)	(−0.74)
GROWTH	−0.001	−0.001	0.001	−0.001	−0.001
	(−0.31)	(−0.35)	(0.61)	(−0.97)	(−0.33)
LEV	−0.076	−0.051	−0.033	0.014	−0.068
	(−2.23)**	(−1.45)	(−1.49)	(0.48)	(−2.03)**
LMVE	0.002	0.005	−0.022	0.007	0.002
	(0.38)	(1.23)	(−6.05)***	(1.56)	(0.43)
PERFORM	−0.041	−0.047	−0.002	−0.030	−0.051
	(−2.17)**	(−1.93)*	(−0.18)	(−2.13)**	(−2.61)
CONTROL	NA	NA	0.021	−0.428	0.003
			(1.19)	(−28.42)***	(0.17)
VCIPO X CONTROL	0.024	−0.005	0.035	−0.040	0.030
	(−2.36)**	(−0.16)	(1.63)	(−1.75)*	(0.98)
Adj. R^2	1.0%	2.0%	2.0%	37.7%	1.0%

***, **, * Denote two-tailed significance at the 1 percent, 5 percent, and 10 percent levels, respectively.
White (1980)-consistent t-statistics are in parentheses.
The sample consists of 2,630 domestic IPO firms going public in the period 1983–2001. The sample firms must have the necessary Compustat data to calculate the abnormal current accruals using the modified Jones model as applied by Teoh et al. (1998b). The following regression is conducted:

$$DCA_i = \alpha_0 + \beta_1 VCIPO_i + \beta_2 AUDQUAL_i + \beta_3 UNDERWRI_i + \beta_4 GROWTH_i + \beta_5 LEV_i + \beta_6 LMVE_i + \beta_7 PERFORM_i + \varepsilon_i$$

where:

DCA_i = abnormal current accruals calculated using (1) the cross-sectional modified Jones (1991) model to obtain the discretionary and nondiscretionary accruals for firm *i* for a given year; and (2) cross-sectional modified Jones (1991) models performance-matched abnormal accrual (the difference between the modified Jones abnormal accrual and that for a matched firm);
$VCIPO_i$ = an indicator variable that equals 1 if the IPO is backed by a VC for firm *i*, and 0 otherwise;
$AUDQUAL_i$ = an indicator variable that equals 1 if the auditor is a Big 6 auditor for firm *i*, and 0 otherwise; this variable is defined to take into account the impact of the various mergers among the leading audit firms in the 1980s and 1990s;
$UNDERWRI_i$ = firm *i*'s underwriter ranking using the Caster-Manaster reputation rankings for IPO underwriters;
$GROWTH_i$ = change in sales between the year of and year prior to the offering for firm *i*, scaled by the sales in the year prior to the offering;
LEV_i = long-term debt scaled by total assets for firm *i*;
$LMVE_i$ = natural logarithm of total assets for firm *i*;

(*continued on next page*)

TABLE 4 (Continued)

$PERFORM_i$ = an indicator variable that equals 1 if the funds flow from operation (Compustat data 110) or cash flow from operations (Compustat data 308) for firm *i* is greater than the two-digit SIC code industry average; industry and year dummies are included in the regressions but results are not tabulated; and

$CONTROL_i$ = the various variables that control for the competing explanations (VC investment, VC exit, presence of a lock-up agreement in the IPO, the Internet bubble years and less aggressive accruals). Investment is the ratio of the dollar amount invested by the VC in firm *i* and the total known dollar amount invested in firm *i* prior to the IPO. Exit equals 1 if the VC sold any proportion of its holdings in the period subsequent to the IPO, and 0 otherwise. Lock-up equals 1 if there is a lock-up agreement present, and 0 otherwise. Bubble equals 1 if the IPO occurred during the period 1997–2000, and 0 otherwise. Less Aggressive equals 1 if the IPO firm's accruals are below the sample median, and 0 otherwise.

performance-matched accruals regressions. We note that the other three VC measures are not significantly associated with abnormal accruals around the IPO. Young, inexperienced VCs appear to have no impact on earnings management, for better or worse. Board representation is also not a significant factor. We rerun the regression with only *VCTOT* interacted with *VCIPO* and report the result in Table 4, column (1). This interaction variable is significantly negative at the 5 percent level, suggesting that the monitoring provided by VCs is associated with lower earnings management.

In additional untabulated univariate tests, we examine the extreme quartiles of abnormal accruals and find that VC investment is significantly higher in the bottom quartile of abnormal accruals compared to the top quartile (t = 2.62). In summary, we provide consistent, strong support for our main research question that VC monitoring is strongly associated with lower earnings management in their IPO firms.

VC Exits

Anecdotal and empirical evidence suggest that VCs exit investments as soon as the lock-up restrictions expire in order to maximize the value of their shares and repatriate the profits to their limited partners (e.g., Casares-Field and Hanka 2001; Healy 2002). We determine IPO firms from which VCs cash out their holdings by examining the TFN Insider Filing Data File. Of the VC-backed IPOs in the sample, VCs cash out from their investments as identified by the TFN Insider Filing Data File in only 1.29 percent (n = 34) of the IPO firms. Of these firms, VCs, on average, sell 23 percent of their pre-IPO shares (median is 10 percent). The mean (median) holding period for these VC firms that subsequently exit from their investment in the IPO firm in the open market is 708 (525) days after the IPO.

We rerun our main tests and include a dummy variable that is equal to 1 if the VC sells any proportion of its holdings in the IPO firm. We then interact this variable with *VCIPO*. The regression result is reported in Table 4, Column (2). The interaction variable is not significant, suggesting that subsequent VC exits from the IPO firm are not associated with earnings management at the time of the IPO. VC presence in the IPO continues to be significantly associated with lower earnings management (t = −2.12).

Internet Bubble

Venture capitalists have been partly blamed for inflating the stock prices of IPO firms during the "Internet bubble" period (e.g., Biddle 2001). Hence, the likelihood of any findings consistent with VCs acting as monitors (and thus mitigating earnings management) during this period is biased downward. We re-estimate Model 1 and control for the "Internet IPO bubble." A dummy variable is included in the regression that is equal to 1 if the IPO took place during 1997–2000 (per, Ritter and Welch 2002; Lee and Wahal 2004), and 0 otherwise. We include this test precisely because it is a period where critics perceive VC monitoring to be at its least effective. If we find that VC presence is still strongly associated with lower levels of earnings management during this period, then we provide additional evidence to support our hypothesis.

The results are presented in Table 4, Column (3). The coefficient on *VCIPO* is negative and significant (t = −2.92), suggesting that VC presence in the non-bubble period is associated with lower earnings management. Both the coefficients on *BUBBLE* and its interaction with *VCIPO* are insignificant, suggesting that: (1) abnormal accruals are not significantly different between the bubble and non-bubble periods; and (2) VC-backed and non-VC-backed IPOs exhibit similar abnormal accruals during the bubble period. Overall, our results are consistent with VCs generally being associated with lower abnormal accruals in the firms that they help bring public. We do not find evidence to support the claims made by critics who believe that VCs opportunistically inflate IPO firm stock prices through earnings management during the Internet bubble period.

Do VCs Always Constrain Earnings Management?

We examine whether there are instances when VCs do not appear to constrain earnings management. We rerun our regression by including a dummy variable that is equal to 1 if the IPO firm's accruals are below the IPO sample median (i.e., these are firms with less aggressive accruals), and 0 otherwise; we also interact this dummy variable with *VCIPO*. Our result is presented in Table 4, Column (4). The interaction variable is significantly negative (t = −1.75), suggesting that VC-backed IPOs have lower accruals among firms with less aggressive use of accruals. *VCIPO* is not significantly negative, suggesting that among IPO firms with more aggressive accruals, VC-backed IPOs do not exhibit significantly different IPO-year accruals than non-VC-backed IPOs. This finding is further explored and discussed below in our tests of the long-run economic consequences of earnings management in VC-backed IPOs.

IPO Lock-Up Provisions

Recent finance literature suggests that IPO lock-up restrictions can have a significant effect on both the investment and exit decisions of various stakeholders and on the market value of the firm itself over time (Brav and Gompers 2003; Ofek and Richardson 2002; Casares-Field and Hanka 2001). We rerun our main tests and examine separately whether (1) the presence of a lock-up provision (reported in Table 4), or (2) the duration of the lock-up agreement are associated with higher accruals at the time of the IPO (not reported for parsimony). Table 4, Column (5) presents the result for the regression that includes a dummy variable that controls for the presence of a lock-up provision and its interaction with *VCIPO*. *VCIPO* continues to be significantly negative in both regressions. Neither lock-up variables nor their interactions with *VCIPO* are significant in the tabulated and untabulated results. This suggests that the presence of a lock-up agreement and the duration of the lock-up agreement are not significantly associated with the IPO-year earnings management behavior of the firm. The evidence on lock-up expirations eliminates the possibility that the results related to VC presence simply reflect a lock-up effect.

Table 5 presents descriptive statistics on the lock-up agreement in our IPO firms. Panel A reports that nearly 60 percent of all IPOs in our research period have a lock-up agreement (62 percent of VC-backed IPOs versus 57 percent of non-VC-backed IPOs). Venture-backed IPOs are more likely to have lock-up provisions than non-VC-backed IPOs (χ^2 = 5.85). The average percentage of shares subject to lock-up provisions is higher in VC-backed IPOs (97 percent) compared to non-VC-backed IPOs (87 percent) (t = 1.94). On average,

TABLE 5
VC- and Non-VC-Backed IPOs and Lock-Up Provisions

Panel A: Lock-Up Provisions and IPO firms

	VC-Backed IPOs	Non-VC-Backed IPOs	Total
Lock-Up Provision	708	843	1,551
No Lock-Up Provision	437	632	1,069
Total	1144	1486	2620
Pearson χ^2			5.8490
p-value			0.016
			t-statistic (z-statistic)[a]
Percentage of Lock-Up Shares[b]	99.6	87.6	1.938*
Duration of Lock-Up Period (days)	190.84	256.43	−9.507***

Panel B: Venture Financing with Lock-Up Provisions and Mean (Median) Discretionary Accruals[c]

	VC-Backed IPOs	Non-VC-Backed IPOs	t-statistic (z-statistic)[a]
Lock-Up Provision	0.030	0.063	−1.610
	(0.027)	(0.043)	(−1.958)*
No Lock-Up Provision	0.018	0.057	−1.495
	(0.014)	(0.051)	(−1.564)
t-statistic	0.752	0.3945	
(z-statistic)	(0.464)	(−0.785)	

Panel C: Venture Financing with Lock-Up Provisions and Mean (Median) Three-Year Long-Run Buy-and-Hold Returns

	VC-Backed IPOs	Non-VC-Backed IPOs	t-statistic (z-statistic)[a]
Lock-Up Provision	−0.146	−0.325	1.519
	(−0.622)	(−0.703)	(1.875)*
No Lock-Up Provision	−0.305	−0.340	0.357
	(−0.613)	(−0.593)	(0.354)
t-statistic	1.261	0.138	
(z-statistic)	(0.535)	(−1.959)*	

***, * Denote two-tailed significance at the 1 percent and 10 percent levels, respectively.

The sample consists of 2,630 domestic IPO firms going public in the period 1983–2001, with available lock-up provision data collected from the Securities Data Corporation Initial Public Offerings Database. Medians are reported in parentheses.

[a] Wilcoxon rank-sum z-statistic.

[b] The percentage of lock-up shares is calculated by the ratio of the number of lock-up shares (as reported in the Securities Data Corporation [SDC] database) to the number of shares outstanding after the IPO. This is consistent with Casares-Field and Hanka (2001). There are 375 (407) VC-backed (non-VC-backed) IPOs with the available information on the number of lock-up shares in the SDC database. The high attrition rate is due to the SDC data being incorrect or missing in 45 percent of the cases (Casares-Field and Hanka 2001).

[c] The performance-matched accruals are reported. Results for the cross-sectional modified Jones discretionary accruals (not tabulated) are qualitatively similar.

the lock-up period for VC-backed IPOs is shorter by approximately 66 days (t = −9.507). Table 4, Panels B and C, report the abnormal accruals and long-run stock returns for VC- and non-VC-backed IPOs with and without lock-up agreements. Overall, this 2 × 2 partitioning provides evidence consistent with the conclusion that the lower abnormal accruals and better long-run stock performance of VC-backed IPOs are not driven by the lock-up agreements. On average, there are no significantly different abnormal accruals and long-run stock returns between VC- and non-VC-backed IPO firms with lock-up agreements (although the median abnormal accruals and stock returns of VC-backed IPOs are slightly lower than their non-VC-backed counterparts).

Is it Really Earnings Management?

Restatements

We have demonstrated that our results are robust across the cross-sectional Jones and performance-matched accruals measures. In this subsection, we examine another proxy for earnings management: earnings restatements. We focus on this alternative proxy, since firms that are forced to restate earnings might be viewed as pushing the bounds of generally accepted accounting principles (GAAP). In addition, the use of restatements avoids some, but not all, problems associated with the use of accruals as surrogates for measuring earnings management (Richardson et al. 2002; Kinney and Libby 2002).

Using the General Accounting Office (GAO) Financial Restatements database, we identify firms in our sample that have restated earnings in the three years subsequent to the IPO because of accounting irregularities or errors. In particular, firms accessing the public equity market have an incentive to overstate earnings in order to increase their share price (Dechow et al. 1996; Dechow and Skinner 2000). As such, we focus on subsequent restatements that lead to a downward revision in earnings. Table 6 presents the results for the sample of IPO firms that restate downward (n = 42). Non-VC-backed IPOs exhibit a higher frequency of earnings restatements subsequent to the IPO that result in a downward revision of earnings compared to VC-backed IPOs ($\chi^2 = 2.75$).

Abnormal Accruals and Post-IPO Performance

If IPO firms use abnormal accruals in the year of the IPO to borrow earnings from future periods, then we would expect these firms to underperform in the post-IPO period

TABLE 6
Venture Capitalist Presence and Restatements

Restatements and IPO firms

	Non-VC-Backed IPO	**VC-Backed IPO**	**Total**
Restatement	29	13	42
No restatement	1456	1132	2588
Total	1458	1145	2630
Pearson χ^2			2.749
p-value			0.097

The sample consists of 2,630 domestic IPO firms going public in the period 1983–2001. Restatement firms are firms that have subsequently restated their earnings downward for reasons of accounting irregularities or errors within three years from the IPO year. The General Accounting Office Financial Restatement database is used to identify the restatements.

(Teoh et al. 1998; Shivakumar 2000). We examine the following measures of firm performance: (1) performance-matched return on sales for the year of the IPO and the three years thereafter; (2) cash flow from operations for the year of the IPO and the three years thereafter; and (3) three-year stock returns. Similar to Teoh et al. (1998), we first match IPO firms to non-IPO firms from the same two-digit SIC code industry that have the closest (to at least 80 percent of) offering-year return on sales as the IPO firm. The performance measure is calculated as the difference between the return on sales (cash flow from operations/sales) of the IPO firm and the matched non-IPO firm.

These results are presented in Table 7, Panel A (performance-matched return on sales) and Panel B (cash flow from operations/sales). In both panels, the average performance of the IPO firm is significantly lower than the matched firm for the two-year period after the issue.[8] Based on return on sales, on average IPO firms underperform their matched non-IPO counterparts by about 17 percent in the first year and 6 percent in the second year; based on cash flow from operations/sales, the underperformance is 23 percent in the first year and 11 percent in the second year. Overall, the evidence is consistent with the notion that IPO firms that report higher earnings in the year of the issue are unable to sustain their firm performance in the post-IPO period.

In further analyses, we match the VC-backed IPO firms to the non-VC-backed IPO firms based on IPO-year ROA and examine the earnings performance in the post-issue period. With both VC and non-VC IPO firms exhibiting similar earnings performance in the year of the issue, and given our expectation that non-VC IPOs exhibit higher abnormal accruals at the time of the IPO holds, then we should observe non-VC-backed IPOs to

TABLE 7

Post-Issue Performance-Matched Return on Sales and Cash Flow from Operations (Scaled by Sales) from the Fiscal Year of the IPO to Three Years After

Panel A: Performance-Matched Differences in Return on Sales (IPO Sales less Non-IPO Sales)

Year	0	1	2	3
Mean	−12.765**	−16.738***	−5.687*	−1.911
Median	0.000	−0.218	−0.437	0.975
Observations	1353	1276	1131	1003

Panel B: Performance-Matched Differences in Cash Flow from Operations/Sales (IPO Cash Flows less Non-IPO Cash Flows)

Year	0	1	2	3
Mean	−14.127**	−23.487**	−10.995*	−4.470
Median	−2.863***	−3.795***	−1.849***	−0.100
Observations	1353	1276	1131	1003

***, **, * Denote two-tailed significance at the 1 percent, 5 percent, and 10 percent levels, respectively.
The sample consists initially of all 2,630 IPOs in the sample. IPO firms are then matched to non-IPO firms in the same two-digit SIC code with the closest (to at least 80% of) offering-year return on sales of the IPO firm. Furthermore, the top and bottom 5 percent of both performance measures are dropped from the sample (consistent with Teoh et al. [1998]). We calculate the performance measure as the difference between return on sales (cash flow from operations/sales) of the IPO firm and matched firm and is reported as a percentage.

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[8] Since firms are matched by return on sales in year 0, the median matched-return on sales in Panel A is zero for year 0.

underperform the VC IPOs in the post-issue period. Figure 2 presents the time series pattern of the earnings performance and abnormal accruals. We observe that the abnormal accruals are higher for non-VC IPOs compared to VC IPOs with similar earnings performance in the year of the IPO. In all three years subsequent to the IPO, these non-VC IPO firms continue to underperform the matched VC IPO firms. Interestingly, the performance of non-VC firms continues to decline between the first and second years after the IPO, whereas we observe the beginnings of performance improvement for VC-backed firms. These findings continue to support our main findings that the presence of a VC in an IPO mitigates earnings management.

Does Earning Management Constraint Explain Superior VC-Backed Long-Run Returns?

Teoh et al. (1998b) show that earnings management at the time of the IPO can explain the long-run stock return underperformance of IPO firms. Other studies show that VC-backed IPOs experience better long-run stock performance than non-VC-backed IPOs (e.g., Brav and Gompers 1997). Consistent with the prior literature, we also show that the long-run stock performance of VC-backed IPO firms is generally superior to the non-VC-backed IPO firms over the three-year window post-IPO (see Figure 3). Table 8 presents the difference in the three-year buy-and-hold returns (BHARs) of VC and non-VC IPOs after we adjust for the endogenous choice of VC financing. We consistently find that VC-backed

FIGURE 2
Post-Issue Abnormal Accruals and Earnings Performance for VC-Backed IPO Firms Matched on IPO-Year ROA and Industry to Non-VC-Backed IPO Firms
(n = 2,630)

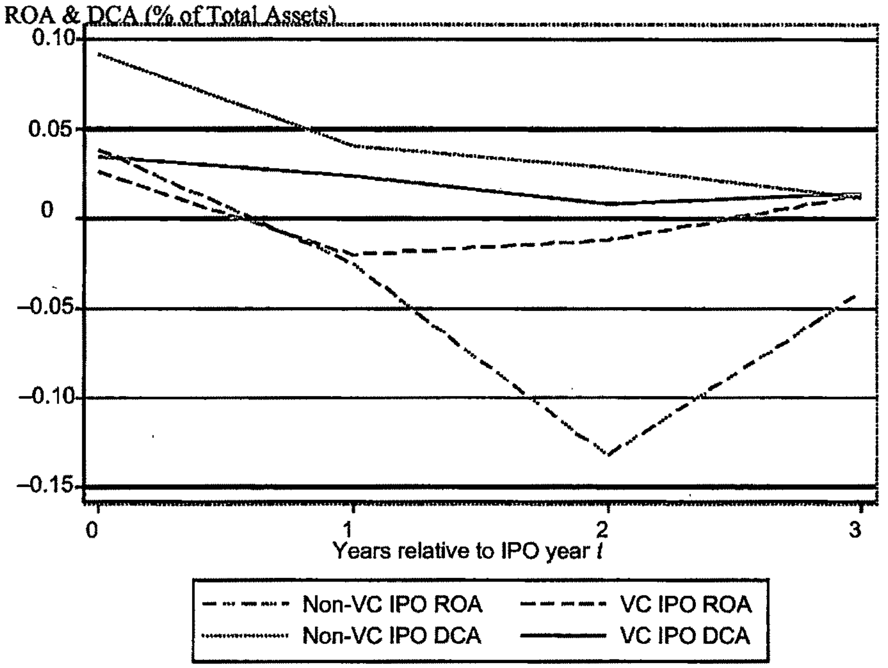
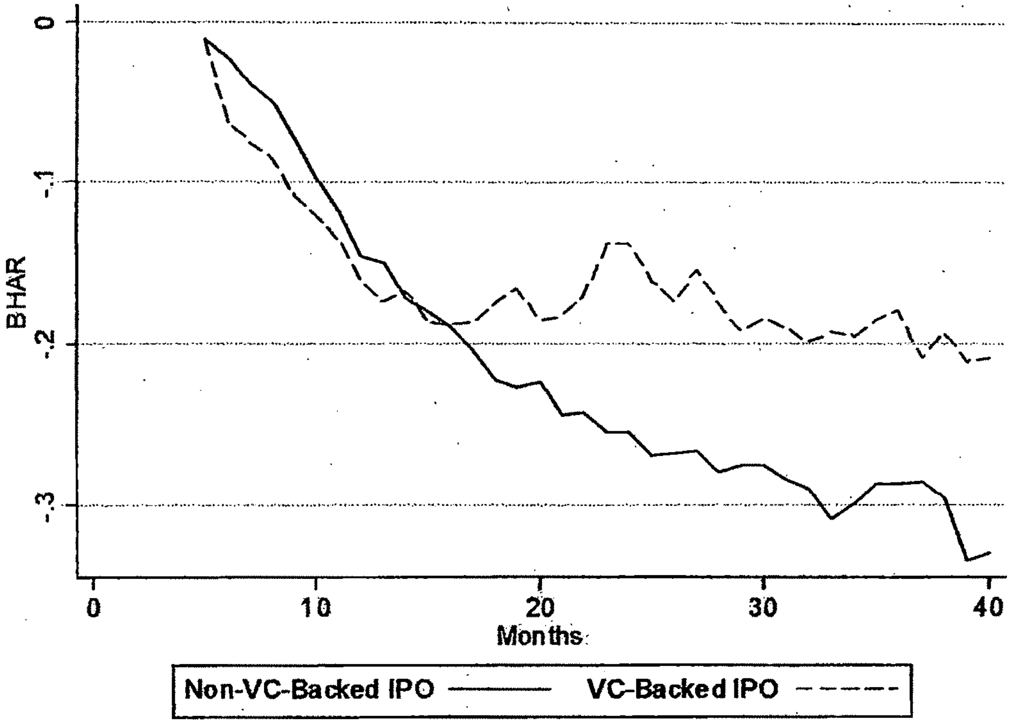


FIGURE 3
Venture and Non-Venture Backed IPOs—Long-Run Post-IPO Stock Returns



The three-year post-issue market-adjusted buy-and-hold returns (BHAR) are calculated four months after the first fiscal year-end (n = 2630).

firms experience higher long-run stock returns regardless of the matching procedure employed. This is consistent with Brav and Gompers (1997) who demonstrate that VC-backed IPOs outperform non-VC-backed IPO firms in the post-IPO period. The average BHARs for the IPO firms in our sample (−27.78 percent) is comparable to the average BHARs in Teoh et al. (1998b) (in the range −15.6 percent to −30.6 percent).

The main thrust of our results thus far consistently suggests that VC presence is associated with lower earnings management at the time of the IPO. However, we find some evidence that VC presence may not always mitigate earnings management as noted above in Table 4, Column (5). If VCs reduce earnings management at the time of the IPO, then does this partially explain why VC-backed IPOs outperform non-VC-backed IPOs in the post-IPO period? Conversely, what do we observe for those IPO firms where VCs appear not to mitigate earnings management? To answer these questions about the economic consequences of VC presence and earnings management, we rank our IPO firms into quartiles based on abnormal accruals in the IPO year and compare the three-year post-IPO buy-and-hold stock returns of the extreme quartiles in Figure 4.

For firms in the quartile with the lowest abnormal accruals (less aggressive firms), we show that the VC-backed firms outperform the non-VC-backed firms over the three-year

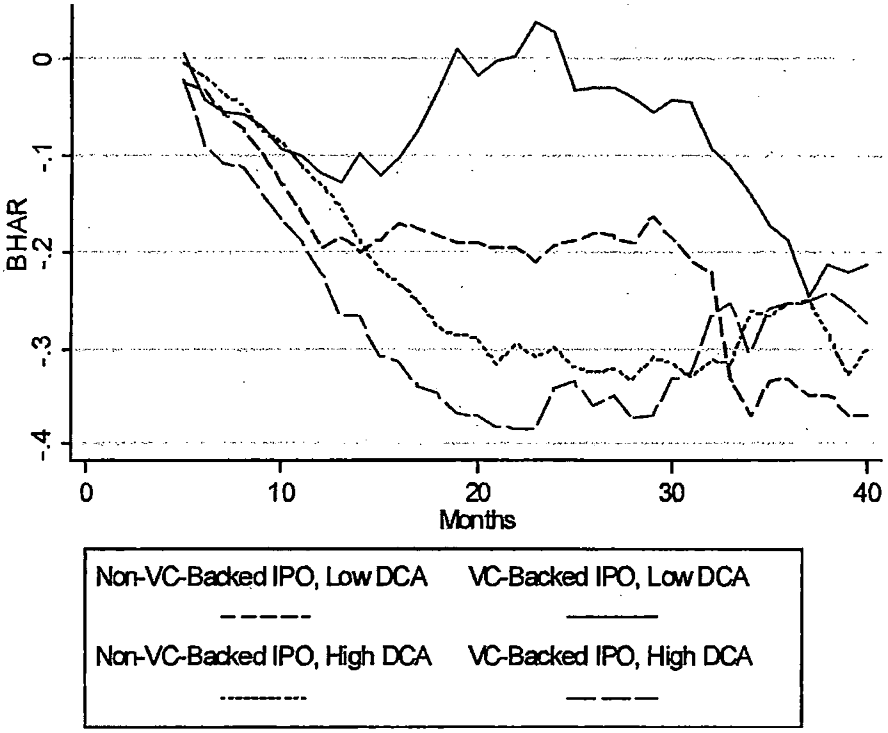
TABLE 8
Venture Capitalist Presence and Long-Run Returns Results

Test of Differences in Long-Run Returns between VC and Matched-Non-VC IPO Firm

	Propensity Score	Gaussian Kernel	Local-Linear Regression Adjusted
Three-year buy-and-hold returns	0.521 (7.931)***	0.153 (2.144)**	0.197 (2.743)***

***, ** Denote two-tailed significance at the 1 percent and 5 percent levels, respectively.
White (1980)-consistent t-statistics are in parentheses.
The sample consists of 2,630 domestic IPO firms going public in the period 1983–2001. The sample firms must have the necessary Compustat data to calculate the abnormal current accruals using the modified Jones model as applied by Teoh et al. (1998). The three-year post-issue market-adjusted buy-and-hold returns are calculated five months after the first fiscal year-end. We adjust for the endogenous choice for VC financing by using two-digit SIC code dummies, headquarter state dummies, year dummies, book value of equity per share, total assets per share, and sales per share as instrumental variables in the first-stage regression. Based on the parameter estimates in the first-stage probit regression, we estimate the probability of VC financing and use this probability measure to match each VC-backed IPO firm to the non-VC-backed IPO firm with the closest probability measure.

FIGURE 4
Long-Run Post-IPO Stock Returns of Venture-Backed and Non-Venture-Backed IPOs in the Top and Bottom Quartiles of Discretionary Current Accruals (DCA)



period post-issuance. However, for firms in the quartile with the highest abnormal accruals (most aggressive firms), we show that the VC-backed firms underperform the non-VC-backed firms. This finding provides new evidence in the literature in two ways. First, we provide some new evidence that suggests that, if VCs mitigate earnings management at the time of the IPO, the VC-backed firms experience better long-run returns compared to their non-VC-backed counterparts. Second, we show that when VCs do not mitigate earnings management, the VC-backed firms underperform their non-VC-backed counterparts.

VII. CONCLUSIONS AND FUTURE RESEARCH

Extant finance theory suggests that venture capitalists (VCs) provide oversight of management activity in their portfolio companies because of their close proximity to the day-to-day operations of the firm. Critics, however, have suggested that the interests of VCs are not aligned with the other stakeholders of the firm and thus do not serve as effective monitors of the firm. We perform a series of tests that consistently show that VC presence and incentives of VCs to monitor are associated with lower earnings management in IPO firms. Our findings also show that VC-backed IPOs outperform non-VC-backed IPOs (in terms of accounting-based performance measures and stock returns) when VCs mitigate earnings. We also find evidence to suggest that when VCs do not mitigate earnings management as much, the IPO firms backed by these VCs tend to experience long-run stock returns underperformance.

Despite a battery of sensitivity tests, the tenor of our main results remains unchanged. We control for the endogenous choice for VC financing, IPO lock-up provisions, and VC exits subsequent to the IPO. Our proxies for earnings management are the cross-sectional modified Jones abnormal accruals and the performance-matched accruals measures, as well as restatements. In addition, we provide evidence that the performance of IPO firms with abnormal accruals declines in the post-issue period. These further tests provide additional comfort that the proxies we use for abnormal accruals capture earnings management.

This study makes significant contributions to both the finance and accounting literatures by providing evidence that VCs behave, on average, as though their incentives are aligned with long-term shareholders with respect to earnings management decisions in an IPO context. We inject empirical evidence into the controversy surrounding whether VCs' interests are aligned or are in conflict with those of the other stakeholders and thus contribute to the debate surrounding these issues.

We also quantify for the first time the value-added impact of VC presence when correlated with earnings management in a sample of IPO firms. We enhance extant accounting research by providing evidence of nonregulatory means for mitigating wealth-reducing earnings management. Finally, we contribute to the accounting literature by documenting the significant corporate governance role of an economically significant, but relatively unstudied, stakeholder—the VC. Our results potentially have implications for current policy-making, but they also indicate rich future research avenues. Future studies may explore in further detail the characteristics of VC firms that appear not to mitigate earnings management. In addition, an open area exists for future research to examine the impact of this key stakeholder on other financial reporting and disclosure settings.

APPENDIX

We employ two approaches to incorporate the endogenous choice for venture capitalist (VC) financing into our research design. For the univariate findings we use a sophisticated matching framework; for the multivariate analyses we use a two-stage switching model. In univariate tests, we are interested in the difference in abnormal accruals between a firm backed by a VC (Y_1) and the *same* firm *without* VC backing (Y_0). However, this is not observable for an individual firm, since the firm either receives VC backing or it does not ($D = 1$ if the firm receives VC financing, and $D = 0$ if the firm does not). To overcome this, we estimate the average effect of VC financing on VC-backed IPOs at the population level conditioned on a vector of firm/industry characteristics ($\mathbf{X}$):

$$E(Y_1 - Y_0|D = 1, \mathbf{X}). \tag{1}$$

Since VC backing is not randomly assigned, but represents an endogenous choice, we cannot simply use the average abnormal accruals of non-VC-backed IPOs.

The bias (B) introduced is indicated as:

$$B(\mathbf{X}) = E(Y_0|D = 1, \mathbf{X}) - E(Y_0|D = 0, \mathbf{X}). \tag{2}$$

Rosenbaum and Rubin (1983) suggest the propensity score method as one means to implement matching while eliminating this bias. Traditional one-to-one matching utilized in recent earnings management research attempts to match two companies on ROA and industry, for example. The propensity score method instead utilizes a vector of variables to calculate to a score for each firm—firms are then matched to each other by this score. The "nearest neighborhood matching" pairs each VC-backed IPO firm to some group of comparable non-VC-backed IPOs (by their respective scores) and then associates the abnormal accruals of the VC-backed firm to the weighted abnormal accruals of its "neighbors" in the comparison group. This weighted average of the scores of more (possibly all) non-VC firms, where the weight given to the individual non-VC firm, is in direct proportion to the closeness of the vector of observables—i.e., "smooth-weighted matching estimators." We employ two approaches toward the calculation of these smooth-weighted matching estimators: (1) kernel-based matching (Gaussian distribution that uses a normal distribution to assign weights to the non-VC-backed IPOs); and (2) local linear regression matching (i.e., only a subsample of the non-VC sample is used to estimate a local regression on the vector of firm/industry characteristics, $\mathbf{X}$).

We approach the endogenous choice in the multivariate regression analyses with a two-step procedure in which the first stage is the estimation of a probit regression that predicts the receipt of venture funding and the second-stage regression uses the estimates from the first stage to provide consistent estimates of the parameters (Madalla 1983). Based on the observed non-randomness of the data in Figure 1 and Table 1, we use two-digit SIC code dummy variables, year dummies, state dummies, book value of equity per share, total assets per share, and sales per share as independent variables in the first-stage regression to determine VC financing.[9] In results (not reported), our first-stage model is significant at levels similar to those of Lee and Wahal (2004). The pseudo-R^2 for our probit model is approximately 18 percent, compared to 19 percent reported by Lee and Wahal (2004).

REFERENCES

Aharony, J., C. J. Lin, and M. P. Loeb. 1993. Initial public offerings, accounting choices, and earnings management. *Contemporary Accounting Research* 10: 61–81.

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[9] We do not use offer characteristics as instrumental variables in the first-stage regression to avoid a look-ahead bias, since most of this information (e.g., net proceeds, offer price) are only known at the time of the IPO.

Baker, M., and P. A. Gompers. 2003. The determinants of board structure at the initial public offering. *Journal of Law and Economics* 46: 569–598.

Barry, C., C. J. Muscarella, J. W. Peavy III, and M. R. Vetsuypens. 1990. The role of venture capital in the creation of public companies. *Journal of Financial Economics* 27: 447–471.

Berle, A. A., and G. C. Means. 1932. *The Modern Corporation and Private Property*. New York, NY: Harcourt, Brace and World.

Biddle, J. 2001. The ones who blew up the balloon. *Washington Post* (January 11): E.1.

Black, B. S., and R. J. Gilson. 1998. Venture capital and the structure of capital markets: Bank versus stock markets. *Journal of Financial Economics* 47: 243–277.

Borrell, J. 2004. Hedge fund regulators set sights on PE funds. *Venture Capital Journal* (September 1): 1.

Brav, A., and P. A. Gompers. 1997. Myth or reality? The long-run underperformance of initial public offerings: Evidence from venture and nonventure capital-backed companies. *Journal of Finance* 52: 1791–1821.

———, and ———. 2003. The role of lock-up provisions in initial public offerings. *Review of Financial Studies* 16: 1–29.

Buckman, R. 2001. Who caused the dot.com crash? Venture capitalists, analysts, entrepreneurs, day traders all play the blame game. *Wall Street Journal* (March 5): B1.

Carter, R. B., F. H. Dark, and A. K. Singh. 1998. Underwriter reputation, initial returns, and the long-run performance of IPO stocks. *Journal of Finance* 53: 285–311.

Casares-Field, L., and G. Hanka. 2001. The expiration of IPO lock-ups. *Journal of Finance* 56: 471–501.

Cohen, D., and N. Langberg. 2005. From private to public: Venture capital financing and the informativeness of earnings. Working paper, University of Southern California, Los Angeles, CA, and University of Houston, Houston, TX.

Darrough, M., and S. Rangan. 2005. Do insiders manipulate earnings when they sell their shares in an initial public offering? *Journal of Accounting Research* 43: 1–33.

Dechow, P. M., R. G. Sloan, and A. P. Sweeney. 1995. Detecting earnings management. *The Accounting Review* 70: 193–226.

———, ———, and ———. 1996. Causes and consequences of earnings manipulation: An analysis of firms subject to enforcement actions by the SEC. *Contemporary Accounting Research* 13: 1–36.

———, and D. J. Skinner. 2000. Earnings management: Reconciling the views of accounting academics, practitioners, and regulators. *Accounting Horizons* 14: 235–250.

Felix, W., C. Tan, and S. Morsfield. 2005a. The demand for auditing in the presence of venture capitalists. Working paper, The University of Arizona, Tucson, AZ, and Baruch College–CUNY, New York, NY.

———, ———, and ———. 2005b. The demand for auditing: Further evidence from initial public offerings. Working paper, The University of Arizona, Tucson, AZ, and Baruch College–CUNY, New York, NY.

Gaver, J. J., and J. S. Paterson. 2001. The association between external monitoring and earnings management in the property-casualty insurance industry. *Journal of Accounting Research* 39: 269–282.

Gompers, P. 1995. Optimal investment, monitoring, and the staging of venture capital. *Journal of Finance* 50: 1461–1490.

———. 1996. Grandstanding in the venture capital industry. *Journal of Financial Economics* 42: 133–156.

———, and J. Lerner. 1997. What drives venture capital fundraising? Working paper, Harvard University, Cambridge, MA, and National Bureau of Economic Research, Cambridge, MA.

———, and ———. 1998a. The determinants of corporate venture capital success: Organizational structure, incentives, and complementarities. Working paper, National Bureau of Economic Research, Cambridge, MA.

———, and ———. 1998b. Venture capital distributions: Short-run and long-run reactions. *Journal of Finance* 53: 2161–2183.

———, and ———. 2002. *The Venture Capital Cycle.* Cambridge, MA: The MIT Press.

Healy, B. 2002. Insider or outsider? Cozy ties with management are raising questions when VCs sit on corporate boards. *Boston Globe* (September 2): D.1.

Heesen, M. 2002. *Comment Re: File No. S7-40-02, Disclosure Required by Section 407 of the Sarbanes-Oxley Act ("Financial Expert").* December 2. Available at: http://www.sec.gov/rules/proposed/s74002/mghessen1.htm.

———. 2003. *Comment Re: File No. SR-NASD-2002-141; Proposed Amendments to NASD Rules 4200 and 4350 Regarding Board Independence and Independent Committees.* April 16. Available at: http://www.sec.gov/rules/sro/nasd2002141/nvca041603.htm.

Hellman, T., and M. Puri. 2002. Venture capital and the professionalization of start-up firms: Empirical evidence. *Journal of Finance* 57: 169–197.

Hribar, P., and D. W. Collins. 2002. Errors in estimating accruals: Implications for empirical research. *Journal of Accounting Research* 40: 105–134.

Hsu, D. H. 2004. What do entrepreneurs pay for venture capital affiliation? *Journal of Finance* 59: 1805–1844.

Jensen, M. C., and W. H. Meckling. 1976. Theory of the firm: Managerial behavior, agency costs and ownership structure. *Journal of Financial Economics* 3: 305–360.

Jones, J. J. 1991. Earnings management during import relief investigations. *Journal of Accounting Research* 29: 193–228.

Kaplan, S. N., and P. Strömberg. 2000. Financial contracting theory meets the real world: An empirical analysis of venture capital contracts. Working paper 7660, National Bureau of Economic Research, Cambridge, MA.

———, and ———. 2001. VCs as principals: Contracting, screening, and monitoring. Working paper 8202, National Bureau of Economic Research, Cambridge, MA.

———, and ———. 2003. Evidence on the venture capital investment process: Contracting, screening and monitoring. In *Venture Capital Contracting and the Valuation of High-Technology Firms,* 73–82. Oxford, U.K. and New York, NY: Oxford University Press.

———, and ———. 2004. Characteristics, contracts and actions: Evidence from venture capitalist analyses. *Journal of Finance* 59: 2177–2210.

Kinney, W. R., and R. Libby. 2002. Discussion of the relation between auditors' fees for nonaudit services and earnings management. *The Accounting Review* 77: 107–114.

Klein, A. 2002. Audit committee, board of director characteristics, and earnings management. *Journal of Accounting and Economics* 33: 375–400.

Kothari, S. P., A. J. Leone, and C. E. Wasley. 2005. Performance matched discretionary accrual measures. *Journal of Accounting and Economics* 39: 163–197.

Lee, P. M., and S. Wahal. 2004. Grandstanding, certification and the underpricing of venture capital backed IPOs. *Journal of Financial Economics* 73: 375–407.

Lerner, J. 1995. Venture capitalists and the oversight of private firms. *Journal of Finance* 50: 301–318.

Madalla, G. S. 1983. *Limited-Dependent and Qualitative Variables in Econometrics.* Econometric Society Monographs No. 3. Cambridge, U.K.: Cambridge University Press.

McNichols, M., and P. Wilson. 1988. Evidence of earnings management from the provision for bad debts. *Journal of Accounting Research* 26 (Supplement): 1–31.

———. 2000. Research design issues in earnings management studies. *Journal of Accounting and Public Policy* 19: 313.

Megginson, W., and K. Weiss. 1991. Venture capitalist certification in initial public offerings. *Journal of Finance* 46: 879–903.

Mills, D. Q. 2001. Who's to blame for the bubble? *Harvard Business Review* 79 (5): 22.

Ofek, E., and M. Richardson. 2002. The valuation and market rationality of Internet stock prices. *Oxford Review of Economic Policy* 18: 265–287.

Richardson, S., I. Tuna, and M. Wu. 2002. Predicting earnings management: The case of earnings restatements. Working paper, University of Pennsylvania, Philadelphia, PA, and Hong Kong University of Science and Technology, Kowloon, Hong Kong.

Ritter, J. R., and I. Welch. 2002. A review of IPO activity, pricing and allocations. *Journal of Finance* 57: 1795–1828.

Rosenbaum, P., and D. Rubin. 1983. The central role of the propensity score in observational studies for causal effects. *Biometrika* 70: 41–55.

Sahlman, W. A. 1990. The structure and governance of venture capital organizations. *Journal of Financial Economics* 27: 473–521.

Shivakumar, L. 2000. Do firms mislead investors by overstating earnings before seasoned equity offerings? *Journal of Accounting and Economics* 29: 339–371.

Stross, R. E. 2000. Venture capitalists aren't villains. *Wall Street Journal* (July 25): A.22.

Subramanyam, K. R. 1996. The pricing of discretionary accruals. *Journal of Accounting and Economics* 22: 249–281.

Teoh, S. H., I. Welch, and T. J. Wong. 1998a. Earnings management and the underperformance of seasoned equity offerings. *Journal of Financial Economics* (50): 63–99.

———, ———, and ———. 1998b. Earnings management and the long-run market performance of initial public offerings. *Journal of Finance* 53: 1935–1974.

———, T. J. Wong, and G. R. Rao. 1998. Are accruals during initial public offerings opportunistic? *Review of Accounting Studies* 2: 175–208.

Tunick, B. 2003. Are VCs next on Fed hit list? Amid secrecy, accounting irregularities and fuzzy IRR calculations flourish in private equity. *The Investment Dealer's Digest* (January): 1.

White, H. 1980. A heteroscedasticity-consistent covariance matrix estimator and a direct test for heteroscedasticity. *Econometrica* 48: 817–838.

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Evidence of the Abnormal Accrual Anomaly Incremental to Operating Cash Flows

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ABSTRACT: Recent research provides evidence that the operating cash flows-to-price ratio subsumes accruals in explaining future annual returns. This suggests that the accrual anomaly is part of the overall value-glamour anomaly and does not represent the mispricing of earnings. We extend the literature by using multiple measures of abnormal accruals and separate analyses of future annual returns and future earnings announcement returns. The results reveal that the operating cash flows-to-price ratio does not subsume abnormal accruals in explaining future annual returns or future announcement returns. We also find that the operating cash flows-to-price ratio does not subsume *total* accruals in explaining future *announcement* returns. These results are not consistent with accruals being a manifestation of the value-glamour anomaly. Our study contributes to the current debate on the existence and the extent of the (abnormal) accrual anomaly. Moreover, the methodology employed can help researchers in exploring mispricing phenomena.

Keywords: *abnormal accruals; market mispricing; operating cash flows-to-price ratio; value-glamour anomaly.*

Data Availability: *All data are from public sources.*

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I. INTRODUCTION

This study tests the relation between future returns and current period abnormal accruals.[1] Previous research documents that total accruals relate negatively to future returns (i.e., the accrual anomaly documented by Sloan [1996]), and the negative relation is due primarily to abnormal accruals (Xie 2001). Desai et al. (2004) provide evidence that the ability of (abnormal) accruals to explain future returns is subsumed by the operating cash flows-to-price ratio (OCF/P). They conclude that the accrual anomaly is part of the overall value-glamour anomaly. We extend this line of research by employing multiple measures of abnormal accruals and testing their relation with both future annual returns and future earnings announcement returns. Using multiple measures of abnormal accruals increases the robustness of conclusions as to whether the abnormal accrual anomaly is part of the overall value-glamour anomaly.[2] Moreover, separately analyzing future annual returns and future announcement returns provides additional insight on whether OCF/P and abnormal accruals represent the same underlying phenomenon. Returns to an earnings-based anomaly are more likely to concentrate around events related to realizations of future earnings (e.g., year-ahead quarterly earnings announcements), whereas returns to a risk-based strategy are more likely to occur evenly over the annual window (Bernard et al. 1997).[3]

We estimate abnormal accruals using the modified Jones model (Dechow et al. 1995).[4] We use different accrual measures, including total and working capital accruals derived from the cash flow statement, and apply different estimation procedures, including industry- and firm-specific regression. We also employ rank models to control for any nonlinearity in the estimation of abnormal accruals. For mispricing tests, we employ univariate and multivariate regression analyses and zero-investment portfolio analyses for both annual returns and announcement returns. For the zero-investment portfolio analysis, we report the mean and standard deviation of abnormal returns across 13 years (1989 to 2001) and document the number of years that have positive abnormal returns. We do this for both annual returns and announcement returns. Furthermore, we test if the 12-day announcement period return (i.e., three-day windows around four quarterly announcements) is significantly greater than random 12-day non-announcement returns for the same firm during the same year.

To control for risk, we follow Sloan (1986), Xie (2001), and Desai et al. (2004) by employing size-adjusted abnormal returns. However, size may not fully control for risk. In our multiple regression tests, we control for additional risk factors. In addition, for our zero-investment portfolio analysis, we calculate returns for randomly generated size-matched portfolios. The procedure involves randomly matching each firm in the long or short position of the abnormal accrual strategy to a firm in the same size decile in the same year. We perform this randomization procedure 100 times and then compare results for abnormal accrual portfolios to results for random size-matched portfolios (see Section II for a more detailed discussion of this procedure).[5]

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[1] Prior research uses the terms "abnormal" or "discretionary" accruals to label the difference between reported and expected accruals. In this paper, we use the term "abnormal" since it is less suggestive as to whether unusual accruals arise from intentional versus unintentional actions of managers.

[2] Desai et al. (2004) use the industry-specific Jones model to estimate abnormal accruals. Xie (2001) uses a total of six different estimates and employs both industry and firm-specific estimates.

[3] Current finance research has not concluded that the value-glamour anomaly, as reflected by variables such as book-to-market, sales growth, and operating cash flows-to-price (OCF/P), is purely related to risk. We find that OCF/P does not relate to future announcement returns, leading us to conclude that OCF/P is not an earnings-based anomaly. However, we do not intend to claim that the OCF/P anomaly is purely risk-based.

[4] Since there is no generally agreed "best" estimate of abnormal accruals in the literature, we initially considered a large number of abnormal accrual models. We considered two classes of accrual expectation models: Jones-based and operating cash flows-based. Our Jones-based models were based on the model derived in Jones (1991) and further modified by Dechow et al. (1995) and Kothari et al. (2005). Our operating cash flows-based models were derived from Dechow and Dichev (2002). Also, as suggested by McNichols (2002), we examined several combinations of the Jones-based models and the operating cash flows-based models. In total, we examined 22 accrual expectation models. For brevity, we report results only for the modified Jones model (Dechow et al. 1995). Interested readers may obtain the full set of results by contacting the authors.

We document that OCF/P does not subsume the abnormal accrual anomaly. Abnormal accruals relate to both future annual returns and future announcement returns, even after controlling for OCF/P. Similar to Desai et al. (2004), we find that OCF/P subsumes total accruals in explaining future *annual* returns. However, we discover that OCF/P does not subsume total accruals in explaining future *announcement* returns. In fact, we find that OCF/P has no explanatory power for future announcement returns. These results make it difficult to conclude convincingly that accruals are simply a manifestation of the value-glamour anomaly. Returns to accrual strategies tend to cluster around future earnings announcements, while returns to the OCF/P strategy occur smoothly over the year. Return behavior of this type is consistent with accruals being more of an earnings-based anomaly and OCF/P being more of a risk-based anomaly.

In the next section, we describe the accrual expectation model and detail the tests of investor mispricing. In Section III, we outline the sample and data, and Section IV reports results. A summary is provided in Section V.

II. RESEARCH DESIGN

Estimation of Abnormal Accruals

To estimate abnormal accruals, we employ the modified Jones model (Dechow et al. 1995):[6]

$$\frac{ACC_{i,t}}{ATA_{i,t}} = \alpha_1 \frac{1}{ATA_{i,t}} + \alpha_2 \frac{(\Delta REV_{i,t} - \Delta REC_{i,t})}{ATA_{i,t}} + \alpha_3 \frac{PPE_{i,t}}{ATA_{i,t}} + \varepsilon_{i,t} \tag{1}$$

where:

ACC = total accruals;
ATA = average total assets;
ΔREV = change in total revenue;
ΔREC = change in accounts receivable; and
PPE = property, plant, and equipment.

Subscripts i and t indicate firm and time, respectively. To reduce measurement error, we employ accrual data as reported on the statement of cash flows (Hribar and Collins 2002). Total accruals = income before extraordinary items (Compustat #18) less operating cash flows (#308 − #124). Equation (1) is estimated within two-digit SIC code each year, and the residuals (ε) estimate abnormal accruals. In our sensitivity tests, we also use working capital accruals for ACC in estimating abnormal accruals. Working capital accruals = increase in accounts receivable + increase in inventory − decrease in accounts payable

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[5] We are indebted to one of our reviewers for proposing the procedures related to random size-matched portfolios and random non-announcement returns. Future research that examines accounting-based market anomalies should consider this approach.

[6] We have analyzed a total of 22 accrual expectation models (see footnote 4) and we find that our main conclusion is robust. We choose to report only the modified Jones model.

− decrease in income taxes payable + increase in other current net assets.[7] We also estimate Equation (1) using rank regressions and firm-specific estimation.

Measuring the Relation between Abnormal Accruals and Future Returns

To examine the nature of the abnormal accrual anomaly, we measure the relation between abnormal accruals and future returns using several tests. Basing conclusions on several tests provides the advantage of not over-relying on a single approach. Tests involving unobservable measures (i.e., abnormal accruals and abnormal returns) are inherently subject to estimation error, and providing results across multiple approaches helps to alleviate some of this uncertainty. In addition, the different approaches employed can provide unique insights and therefore help in making stronger conclusions.

Our first test of investor mispricing involves a simple regression of returns in year $t+1$ on the decile rank of abnormal accruals in year t:[8]

$$Return_{t+1} = \beta_0 + \beta_1(-Rank\ of\ Abnormal\ Accruals_t) + \varepsilon_{t+1}. \qquad \text{(Test 1)}$$

In each year, abnormal accruals are ranked evenly into deciles from 0 to 9. These ranks are divided by 9 to create a variable ranging from 0 to 1. We multiply the decile rank by -1 so that β_1 is expected to be positive. The coefficient is interpreted as the return to a zero-investment portfolio, *assuming a linear relation across the deciles.* A long (short) position is taken in firms with low (high) abnormal accruals. Thus, more positive values of β_1 imply greater mispricing identified by the estimation procedure.

For our second test, we consider the extent to which abnormal accruals relate to future returns, *after controlling for other variables.* If mispricing relates to abnormal accruals, then abnormal accruals should remain significantly related to future returns in the presence of other variables. We estimate the following multivariate model:

$$\begin{aligned} Return_{t+1} = {} & \beta_0 + \beta_1(-Rank\ of\ Abnormal\ Accruals_t) \\ & + \beta_n(Rank\ of\ Control_{n,t}) + \varepsilon_{t+1}. \end{aligned} \qquad \text{(Test 2)}$$

In our tests, we include three control variables: (1) OCF/P, (2) book-to-market ratio, and (3) sales growth. Each of these variables has been shown by prior research to explain future abnormal returns (e.g., Lakonishok et al. 1994; Desai et al. 2004).[9] Firms with higher OCF/P, higher book-to-market ratios, and lower sales growth are labeled "value" stocks and have historically outperformed "glamour" stocks. Researchers differ on whether these findings relate to investor mispricing (e.g., Lakonishok et al. 1994; La Porta et al. 1997; Dechow and Sloan 1997) or compensation for risk (Fama and French 1992, 1996; Doukas et al. 2002). We are particularly interested in controlling for OCF/P. Desai et al. (2004) report that the total accrual anomaly persists after controlling for the book-to-market ratio and sales growth, but not after controlling for OCF/P. The purpose of our paper is to determine whether abnormal accruals are part of the overall value-glamour anomaly as well.

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[7] That is, working capital accruals = −(#302 + #303 + #304 + #305 + #307). Compustat reports the indirect method adjustments from the statement of cash flows. Hence, a negative sign is needed for generating accruals.

[8] As discussed later, we measure annual returns as the 12-month buy-and-hold size-adjusted return beginning three months after the fiscal year-end. We measure announcement period returns over the four three-day periods surrounding next year's quarterly earnings announcements (i.e., a 12-day interval).

[9] In additional tests, we also control for market beta. Results when including this additional control variable are similar to those reported.

Our third test involves the usual calculation of the return to a zero-investment portfolio. We take a long position in firms in the most negative abnormal accrual decile in year t and a short position in firms in the most positive abnormal accrual decile in year t:

$$\textit{Portfolio Return}_{t+1} = \textit{Low Decile Return}_{t+1} - \textit{High Decile Return}_{t+1}. \qquad \text{(Test 3)}$$

The more positive the portfolio return, the greater the evidence of mispricing. In interpreting the results of Tests 1–3, an important caveat is that a significant relation with future abnormal returns may *not* be evidence of investor mispricing. The abnormal accrual estimates examined in this study may rank firms according to cross-sectional differences in risk, instead of identifying mispriced securities. Since firms with greater risk are expected to earn a higher rate of return, a relation between future returns and current accounting information could be found in the absence of investor mispricing. While we attempt to control for cross-sectional differences in risk by using size-adjusted returns in our analyses and explicitly controlling for other variables (i.e., Test 2), it is always the case that "other" risk factors may provide the explanation. We are interested in identifying abnormal accruals that appear to relate to investors' inability to accurately understand the implication of current earnings for future earnings, causing securities to be temporarily mispriced. Therefore, we deem it important to determine whether results occur because of risk-based explanations or investor mispricing. We test this in three ways: (1) examine the relation between current variables and future earnings *announcement* returns, (2) test if abnormal returns are generated consistently across years, and (3) compare actual zero-investment portfolio returns with returns from random size-matched portfolios.

The first two tests are proposed by Bernard et al. (1997). First, they suggest that if anomalous results occur due to investor mispricing and this mispricing relates to an earnings-based anomaly, then abnormal returns should concentrate around future earnings announcement dates. It is the release of future earnings when investors are more likely to realize any prior mispricing related to earnings, and prices will correct. Accordingly, we extend Tests 1, 2, and 3 to announcement returns. For Test 3, we note the percentage of the portfolio's annual return occurring during earnings announcements. The announcement interval consists of the four three-day periods surrounding next year's quarterly earnings announcements (i.e., 12 days), which comprises approximately 5 percent of the annual interval. According to Bernard et al. (1997), if the relation with returns in year $t+1$ is the result of an earnings-based anomaly, then the portion of abnormal returns occurring at the earnings announcement will likely be greater than 5 percent. If the anomalous results are risk-based, then abnormal returns are more likely to occur evenly throughout the period and not concentrate around earnings announcement dates.[10] We extend Bernard et al.'s (1997) announcement return test by also comparing announcement returns with randomly generated 12-day non-announcement returns for the same firm in the same year. We refer to this as the excess announcement returns.

The second test proposed by Bernard et al. (1997) relates to the consistency in portfolio returns. They suggest that anomalous returns from a zero-investment portfolio are more representative of investor mispricing when the annual return is consistently positive. A zero-investment portfolio that produces a positive return because of cross-sectional differences in risk will show variability in annual returns (i.e., some years will be positive while others will be negative). Therefore, we also examine the number of years that the zero-investment portfolio produces positive returns. We do this test using annual returns (as suggested by Bernard et al. [1997]) and using announcement period returns:

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[10] Bernard et al. (1997) attempt to distinguish between market mispricing and compensation for risk for six anomalies identified in the literature. Two of the variables relate to the value-glamour anomaly: book-to-market ratio and price-to-earnings ratio. They find that announcement returns to the book-to-market ratio are opposite to that predicted by unexpected changes in earnings, and announcement returns to the earnings-to-price ratio are negligible. The fact that both variables produce significantly positive annual returns but do not have significant announcement returns in the predicted direction leads them to conclude that both variables "are more likely to represent risk premia" (Bernard et al. 1997, 89). They do not evaluate OCF/P.

$$\textit{Years Portfolio Positive}_{t+1} = \textit{Number of Positive Portfolio Returns}_{t+1}. \qquad \text{(Test 4)}$$

Our final test for determining whether anomalous returns relate to cross-sectional differences in risk or investor mispricing is to examine returns to random size-matched portfolios. While the returns employed in our analyses are size-adjusted, this may not sufficiently control for risk. Generating returns from random size-matched portfolios provides evidence of whether returns to the abnormal accrual strategy are what one might expect from a risk-based strategy. This randomization procedure involves randomly matching each firm in the long or short position to a firm in the same size decile in the same year. We then subtract the average return of the random stocks in the short position from the average return of the random stocks in the long position. We do this each year and calculate the mean of the annual zero-investment portfolio returns over our 13-year sample period. We perform this randomization procedure 100 times and then compare results for abnormal accrual portfolios to results for random size-matched portfolios for Tests 3 and 4. If returns to abnormal accrual portfolios and to random size-matched portfolios are similar, then this suggests that it is the size-related risk component that drives returns to the abnormal accrual anomaly. In addition to comparing means and number of years with positive returns, we also compare standard deviations in returns between the abnormal accrual portfolios and random size-matched portfolios. Standard deviations provide additional evidence of the risk nature (i.e., variance) of portfolio returns.

III. DATA AND SAMPLE PROCEDURES

Our sample includes all firm-year observations that have the necessary financial data from Compustat and stock return data from CRSP to generate our mispricing tests over the 1989–2001 period, excluding firms with SIC codes 6000–6999, firms with negative book values, or firms with sales less than 1 million.[11] Year-ahead abnormal returns are calculated as the 12-month buy-and-hold size-adjusted return beginning three months after the year-end.[12] We winsorize size-adjusted returns greater than 225 percent, as large returns can cause misleading inferences in tests of mispricing (Kraft et al. 2006). Approximately 1 percent of our observations meet this criterion.

The control variables used in our study are OCF/P, book-to-market ratio, and sales growth. OCF/P is operating cash flows reported from the statement of cash flows over the market value at the end of the third month following the fiscal year-end; book-to-market ratio is the ratio of the fiscal year-end book value of equity to the market value at the end of the third month after fiscal year-end; sales growth is the average of annual growth in sales over the previous three years. The final sample size for our primary test is 25,018 firm/year observations over the period 1989–2001.

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[11] Our sample selection is similar to Desai et al. (2004) except that we also restrict our observations to have data from the statement of cash flows.

[12] We use Eventus to generate size-adjusted return. Eventus uses the equally weighted method to calculate size-portfolio returns. Many studies calculate returns beginning three months after fiscal year-end. Our results are similar when we change the return calculation period to begin four months after the fiscal year-end. Based on an untabulated analysis, we find that calculating returns three months after the year-end (instead of four months) results in a higher correlation with the return extending from the first quarter's earnings announcement through the fourth quarter's earnings announcement. We choose to use the return interval beginning three months after the fiscal year-end to better align with the year-to-year earnings announcement return.

Panel A in Table 1 reports descriptive statistics. Annual abnormal returns and announcement abnormal returns have similar means (0.025 and 0.024, respectively), but the variance is much higher for annual returns. The higher variance is expected given that announcement returns reflect primarily accounting information, while non-announcement returns are also affected by cross-sectional differences in risk. Total accruals have a negative mean and median (−0.054 and −0.046) and working capital accruals have a positive mean and median (0.023 and 0.015). This occurs primarily because total accruals include depreciation. Total accruals also include special items, which are negative on average. OCF/P

TABLE 1
Descriptive Statistics[a]

Panel A: Distributions

Variables[b]	**Mean**	**Standard Deviation**	**Q1**	**Median**	**Q3**
Annual Abnormal Returns$_{t+1}$	0.025	1.078	−0.422	−0.120	0.215
Annc. Abnormal Returns$_{t+1}$	0.024	0.198	−0.080	0.004	0.103
Total Accruals$_t$	−0.054	0.153	−0.100	−0.046	0.006
Working Capital Accruals$_t$	0.023	0.096	−0.018	0.015	0.060
Operating Cash Flows$_t$*/Price*$_t$	0.094	3.146	−0.005	0.060	0.143
Abnormal Accruals$_t$	0.000	0.389	−0.047	0.009	0.071

Panel B: Correlations[c]

Variables	*Annual Abnormal Returns*$_{t+1}$	*Annc. Abnormal Returns*$_{t+1}$	*Total Accruals*$_t$	*Working Capital Accruals*$_t$	*Operating Cash Flows*$_t$*/Price*$_t$	*Abnormal Accruals*$_t$
Annual Abnormal Returns$_{t+1}$		0.279	−0.052	−0.051	−0.006#	−0.039
Annc. Abnormal Returns$_{t+1}$	0.308		−0.063	−0.047	0.031#	−0.056
Total Accruals$_t$	−0.052	−0.050		0.692	−0.212	0.663
Working Capital Accruals$_t$	−0.071	−0.033	0.772		−0.291	0.476
Operating Cash Flows$_t$*/Price*$_t$	0.155	0.039	−0.417	−0.456		−0.166
Abnormal Accruals$_t$	−0.068	−0.047	0.747	0.619	−0.331	

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[a] The sample period is 1989–2001.

[b] *Annual Abnormal Returns* = size-adjusted returns over the 12-month period beginning three months after the fiscal year-end; *Annc. Abnormal Returns* = 12-day size-adjusted returns, consisting of the four three-day periods surrounding quarterly earnings announcements in year $t+1$. Accruals and operating cash flows are amounts from the statement of cash flows; *Total accruals* = income before extraordinary items (Compustat #18) less operating cash flows (#308 − #124); *Working Capital Accruals* = −(#302 + #303 + #304 + #305 + #307). Abnormal Accruals are residuals from the modified Jones model (Dechow et al. 1995), estimated within two-digit SIC code each year.

[c] The upper-right corner represents the average Pearson correlation coefficient across sample years. The lower-left corner represents average Spearman correlation coefficients. All correlation coefficients are significantly different from zero at the 0.05 level, except those with a "#," which are not significant.

has a positive mean and median (0.094 and 0.060). Abnormal accruals from the modified Jones model have a mean of zero and a slightly positive median (0.009).

Panel B of Table 1 reports the average of yearly correlation coefficients. Pearson correlation coefficients are reported on the upper-right corner and Spearman correlation coefficients are reported on the lower-left corner. The Spearman correlation coefficient between annual returns and announcement returns is 30.8 percent. Both returns are negatively related with total accruals, working capital accruals, and abnormal accruals and positively related with OCF/P. All of the accrual measures (i.e., total accruals, working capital accruals, and abnormal accruals) are positively correlated, and they are all negatively correlated with OCF/P. The Spearman correlation coefficients are all significant and they are higher in magnitude than their corresponding Pearson correlation coefficients. Interestingly, the Pearson coefficients for the OCF/P and returns variables are not significant. This confirms the importance of using ranks of OCF/P in our regression analysis reported later in the paper.

IV. RESULTS

Tests of the Accrual Anomaly

In this section we first provide results of the accrual anomaly to establish consistency with prior research. Sloan (1996) concludes that investors overestimate the persistence of total accruals, leading to a negative relation between accruals and future abnormal returns.[13] In Panel A of Table 2, we report tests using annual returns, as is commonly done in the literature. In Panel B, we replicate all of our tests using announcement returns. For comparison, we provide analyses for both total accruals and working capital accruals and all inferences are the same between the two.

The univariate regression coefficient is 0.079 for total accruals and 0.093 for working capital accruals, each of which is significant. Recall that we multiply the rank of accruals by −1 so that the expected relation is positive. Similar to previous findings, firms with low accruals have a higher price performance in the following year than do firms with high accruals. The annual returns to the zero-investment portfolio (0.090 and 0.091 for total and working capital accruals, respectively) are similar in magnitude to the regression coefficients and are significant. The standard deviations of the annual returns to the zero-investment portfolios for total accruals and working capital accruals are 0.086 and 0.093, respectively.

As discussed before, we also report (in parentheses) average abnormal returns and their standard deviation for 100 randomly generated size-matched portfolios. The average annual returns to the random zero-investment portfolios are negative and close to zero (−0.020 and −0.021). This confirms that significantly positive returns of the accrual strategy represent anomalous market behavior. The average standard deviations for the random portfolios (0.072 and 0.071) are slightly smaller in magnitude than the standard deviations for the accrual portfolios, but these differences are small compared to the differences in average returns. Overall, comparisons to the random portfolios strengthen our conclusion that the significant relation between current year accruals and future annual returns occurs because of accruals.

We show the number of years that annual returns to the zero-investment portfolio are positive and the average number of years that annual returns to the 100 randomly generated

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[13] Zach (2003) provides a number of alternative measurement and sampling approaches to determine the robustness of the accrual anomaly. He finds that certain factors (e.g., mergers, divestitures, abnormal return calculation, return interval, and non-NYSE/AMEX firms) play a part in explaining the accrual anomaly. However, he concludes that a substantial portion of the accrual anomaly remains unexplained.

TABLE 2
Relation between Accruals and Future Returns[a]

	Regressions[b]		Zero-Investment Portfolios[c]			
	Coef.	Coef. with Controls	Mean	Standard Deviation	Years Positive	Excess
Panel A: Annual Returns[d]						
Total Accruals	0.079*	0.009	0.090* (−0.020)	0.086 (0.072)	12 (6)	
Working Capital Accruals	0.093*	0.014	0.091* (−0.021)	0.093 (0.071)	12 (6)	
Operating Cash Flows/Price	0.205*		0.176* (0.020)	0.168 (0.069)	12 (8)	
Panel B: Announcement Returns[e]						
Total Accruals	0.036*	0.034*	0.058* (0.002)	0.020 (0.031)	12 (7)	0.056*
Working Capital Accruals	0.024*	0.019*	0.051* (0.003)	0.028 (0.032)	12 (7)	0.047*
Operating Cash Flows/Price	0.015^		0.017^ (−0.003)	0.044 (0.035)	6 (6)	−0.008

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*, ^ Significant at the 0.05 and 0.10 levels, respectively, using a one-tailed t-test that the mean is greater than zero.

[a] The sample period is 1989–2001. Accruals and operating cash flows are amounts from the statement of cash flows.

[b] "Coef." represents the average annual coefficient in a regression of future size-adjusted returns on decile ranks of the variable. Decile ranks are 0 to 9, scaled by 9. For total accruals and working capital accruals, decile ranks are multiplied by −1. "Coef. with Controls" represents the average annual coefficient in a regression of future size-adjusted returns on decile ranks of the variable when decile ranks of operating cash flows-to-price ratio, book-to-market ratio, and sales growth are included in the model.

[c] Zero-investment portfolio returns in year $t+1$ are calculated by subtracting the average size-adjusted return of firms in the highest abnormal accruals decile (i.e., short position) from the average size-adjusted return of firms in the lowest abnormal accruals decile (i.e., long position). For operating cash flows-to-price, the long (short) position is the highest (lowest) decile. The "Mean" amount is the average return to the portfolio over the sample period. The "Standard Deviation" is the standard deviation of the portfolio's returns. "Years Positive" represents the number of years (out of a possible 13) that the zero-investment portfolio produces a positive return in year $t+1$. "Excess" is calculated by taking the announcement return of each firm in the zero-investment portfolio and subtracting a random 12-day non-announcement return for the same firm in the same year. Amounts in parentheses represent averages for 100 random size-matched zero-investment portfolios (see Section II for a description of the randomization procedure).

[d] Annual returns are measured as size-adjusted returns over the 12-month period beginning three months after the fiscal year-end.

[e] Announcement returns are calculated as the 12-day size-adjusted return, consisting of the four three-day periods surrounding quarterly earnings announcements in year $t+1$.

size-matched portfolios are positive. If investors consistently misprice securities, then returns should be positive in all years. If the accrual anomaly reflects risk, then the sign of returns should show variation across years. Comparison to the number of years that the random size-matched portfolios produce positive returns provides some expectation for assessing the anomalous behavior of accruals. Based on random size-matched portfolios, the number of positive years is six (out of 13), which is much smaller than the number of positive years for the total accrual portfolio (12) or the working capital accrual portfolio (12). These results provide evidence consistent with abnormal returns to the accrual strategy not being risk-based.

We are primarily interested in whether the accrual anomaly remains once controlling for OCF/P, book-to-market ratio, and sales growth. As shown in the second column of results, the coefficients on accruals (total and working capital) are no longer significant when adding control variables. Untabulated results reveal that it is OCF/P, rather than the book-to-market ratio or sales growth, that subsumes the accrual anomaly. Consistent with results in Desai et al. (2004), our evidence suggests that the ability of accruals to explain future annual returns is subsumed by OCF/P.

In Panel B of Table 2, we turn to analyses for announcement period returns. The announcement period consists of the four three-day periods surrounding the quarterly earnings announcements. Both total accruals and working capital accruals are significantly related to future announcement returns. The coefficients on total accruals and working capital accruals are 0.036 and 0.024, respectively. The returns to zero-investment portfolios are also significant and of similar magnitude. Approximately 64.4 percent (= 0.058/0.090) of the annual abnormal return of the zero-investment portfolio occurs during earnings announcements, which comprise approximately 5 percent of the annual interval (12/250 = 5 percent of the trading days). The mean announcement return for the random portfolios is very close to zero. Moreover, we find that excess announcement returns, measured as the actual announcement return minus the random non-announcement return for each firm in the same year, are similar to the actual announcement returns for both total and working capital accruals. Random non-announcement 12-day abnormal returns to the zero-investment portfolio are close to zero, while returns during earnings announcements appear to be "large." The fact that most of the annual return to the accrual anomaly is concentrated during events related to the realization of future earnings suggests that investors misprice current year accruals and correct their beliefs in the following year as future earnings become known (Bernard et al. 1997). The standard deviations in announcement returns are 0.020 and 0.028 for total accruals and working capital accruals, respectively, which are lower than the average standard deviations for the random portfolios (0.031 and 0.032, respectively). These results also suggest that returns to the accrual portfolio are not size-related and are less likely to represent cross-sectional differences in risk.

Most importantly for our study is that accruals continue to relate to future announcement returns after controlling for OCF/P. As shown in the first two columns, the coefficient on accruals is significant and of similar magnitude whether controlling for OCF/P or not.[14] The coefficient on total accruals (working capital accruals) is 0.036 (0.024) in the univariate regression, while the coefficient is 0.034 (0.019) after controlling for OCF/P. In addition, untabulated results show that OCF/P is not significantly related to announcement period returns. While OCF/P subsumes accruals when explaining future annual returns, the opposite is true when explaining future announcement returns.

Finally, we provide results for the OCF/P strategy in Table 2. The analyses are similar to those reported for accruals with the portfolio ranks based on the ascending (instead of descending) order of OCF/P. There are two noticeable differences in results. First, the zero-investment portfolio returns are approximately twice that of the accrual anomaly (0.176), but the announcement period returns are only one-third that of the accrual anomaly (0.017). While approximately 9.7 percent of the zero-investment portfolio returns to OCF/P occurs during earnings announcements, it is much lower than the 64.4 percent for the accrual strategy. After controlling for non-announcement returns, we find that excess announcement returns are no longer significant. This implies that OCF/P will not subsume the ability of accruals to explain future announcement returns, as shown by the multiple regression analysis discussed previously.

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[14] Note that we have also included other control variables such as the book-to-market ratio and sales growth.

The second difference in results relates to the standard deviation of returns. For OCF/P, the standard deviation of annual returns and the standard deviation of announcement returns are substantially higher than their random size-matched portfolios. This is not the case for accruals. Similarly, the standard deviation of returns from the OCF/P portfolio is about twice that of accrual portfolios (0.168 versus 0.086 and 0.093 for annual returns and 0.044 versus 0.020 and 0.028 for announcement returns). The results for standard deviation of returns imply that the OCF/P anomaly reflects more of a risk-based anomaly than does the accrual anomaly. Portfolios constructed using OCF/P produce positive (and volatile) returns compared to random size-matched portfolios.

In summary, the results in Table 2 reveal that the return behavior of OCF/P is different than that of accruals. OCF/P behaves more like a risk-based proxy than do accruals. Returns to the OCF/P strategy are positive and volatile and do not cluster around future earnings announcements. On the other hand, returns to the accrual strategy appear to be more earnings-based, as they cluster around future earnings announcements. However, it is still noted that OCF/P subsumes total accruals in explaining future annual returns. Next, we repeat these tests for abnormal accruals, our variable of primary interest.

Test of the Abnormal Accrual Anomaly

We report test of the abnormal accrual anomaly in Table 3. Similar to Table 2, we report results for our regression tests and also compare results between the abnormal accrual portfolio and the 100 random size-matched portfolios (in parentheses). As shown in the first column, the coefficient on abnormal accruals in a univariate regression with future returns is positive and significant. The coefficient is similar in magnitude to the zero-investment portfolio return. Of primary interest to our study, the coefficient on abnormal accruals remains positive and significant after controlling for OCF/P. In other words, OCF/P does not subsume abnormal accruals in explaining future annual returns.[15] This conclusion is different from that of total accruals in Table 2.

Panel B of Table 3 shows that the inability of OCF/P to subsume abnormal accruals is especially apparent for announcement returns. In the univariate regression, the coefficient on abnormal accruals is 0.030 and significant. After controlling for OCF/P (and the book-to-market ratio and sales growth), the coefficient on abnormal accruals reduces to only 0.024 and easily remains significant. The mean return to the zero-investment portfolio is quite high (0.050) for a 12-day interval, and returns do not appear risk-related, as the mean returns to the size-matched portfolios is zero. The standard deviation of the announcement returns to the zero-investment portfolio is approximately one-half that of the random portfolios, and announcement returns to the abnormal accrual strategy are positive in 12 out of 13 years. Moreover, the excess announcement return is significantly positive.

The results in Table 3 provide the conclusion that abnormal accruals are incremental to OCF/P in explaining future returns. Abnormal accruals are not a simple manifestation of the value-glamour anomaly as captured by OCF/P, the book-to-market ratio, and sales growth. This conclusion is especially apparent by noting the clustering of annual returns

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[15] Desai et al. (2004) also conduct a multivariate regression test of annual abnormal returns on abnormal accruals estimated using the industry-specific Jones model. Recall that we present results using the industry-specific *modified* Jones model. They report a t-statistic of 1.74, which is insignificant at the 5 percent level. When we estimate abnormal accruals using the Jones model, our t-statistic is 1.98, which is significant at the 5 percent level. The relatively small difference between our results and theirs is likely to be sample specific. Desai et al. (2004) do not test announcement period abnormal returns.

TABLE 3
Relation between Abnormal Accruals and Future Returns[a]

Regressions[b]		Zero-Investment Portfolios[c]			
Coef.	**Coef. with Controls**	**Mean**	**Standard Deviation**	**Years Positive**	**Excess**
Panel A: Annual Returns[d]					
0.103*	0.051*	0.098* (−0.001)	0.097 (0.069)	11 (6)	
Panel B: Announcement Returns[e]					
0.030*	0.024*	0.050* (0.000)	0.016 (0.030)	12 (7)	0.048*

* Significant at the 0.05 level using a one-tailed t-test that the mean is greater than zero.

[a] The sample period is 1989–2001. Accruals and operating cash flows are amounts from the statement of cash flows. Abnormal accruals are estimated using the modified Jones model within each two-digit SIC code each year.

[b] "Coef." represents the average annual coefficient in a regression of future size-adjusted returns on decile ranks of abnormal accruals. Decile ranks are 0 to 9, scaled by 9. Decile ranks are multiplied by −1. "Coef. with Controls" represents the average annual coefficient in a regression of future size-adjusted returns on decile ranks of abnormal accruals when decile ranks of operating cash flows-to-price ratio, book-to-market ratio, and sales growth are included in the model.

[c] Zero-investment portfolio returns in year $t+1$ are calculated by subtracting the average size-adjusted return of firms in the highest abnormal accruals decile (i.e., short position) from the average size-adjusted return of firms in the lowest abnormal accruals decile (i.e., long position). The "Mean" amount is the average return to the portfolio over the sample period. The "Standard Deviation" is the standard deviation of the portfolio's returns. "Years Positive" represents the number of years (out of a possible 13) that the zero-investment portfolio produces a positive return in year $t+1$. "Excess" is calculated by taking the announcement return of each firm in the zero-investment portfolio and subtracting a random 12-day non-announcement return for the same firm in the same year. Amounts in parentheses represent averages for 100 random size-matched zero-investment portfolios (see Section II for a description of the randomization procedure).

[d] Annual returns are measured as size-adjusted returns over the 12-month period beginning three months after the fiscal year-end.

[e] Announcement returns are calculated as the 12-day size-adjusted return, consisting of the four three-day periods surrounding quarterly earnings announcements in year $t+1$.

to the abnormal accruals strategy during earnings announcements. Annual returns to the OCF/P strategy do not cluster around future earnings announcements. In untabulated results we find that OCF/P is not incrementally significant in explaining future announcement returns, once controlling for abnormal accruals.

Abnormal Accruals Estimation Using Working Capital Accruals

In the previous section, we estimate abnormal accruals as a portion of total accruals. As a sensitivity test, in this section we estimate abnormal accruals as a portion of working capital accruals. These results are reported in Table 4. The overall tenor of the results is similar to that in Table 3. In univariate tests, abnormal accruals have a significant relation with future abnormal returns. After controlling for OCF/P, evidence of the abnormal accrual anomaly remains for both annual returns and announcement period returns.

Abnormal Accruals Estimation Using Rank Models

In this section, we consider the use of rank models to estimate abnormal accruals. Recall that the results in Tables 3 and 4 are based on the traditional approach to estimating abnormal accruals, which assumes a linear relation (i.e., OLS regression) between accruals and financial variables. Those results are reported using raw measures in the first stage

TABLE 4
Relation between Abnormal Accruals and Future Returns Using Working Capital Accruals to Estimate Abnormal Accruals[a]

Regressions[b]		Zero-Investment Portfolios[c]			
Coef.	Coef. with Controls	Mean	Standard Deviation	Years Positive	Excess
Panel A: Annual Returns[d]					
0.118*	0.052*	0.086* (−0.010)	0.087 (0.069)	11 (6)	
Panel B: Announcement Returns[e]					
0.027*	0.023*	0.046* (0.001)	0.023 (0.034)	11 (7)	0.041*

* Significant at the 0.05 level using a one-tailed t-test that the mean is greater than zero.

[a] The sample period is 1989–2001. Accruals and operating cash flows are amounts from the statement of cash flows. Abnormal accruals are estimated using the modified Jones model within each two-digit SIC code each year.

[b] "Coef." represents the average annual coefficient in a regression of future size-adjusted returns on decile ranks of abnormal accruals. Decile ranks are 0 to 9, scaled by 9. Decile ranks are multiplied by −1. "Coef. with Controls" represents the average annual coefficient in a regression of future size-adjusted returns on decile ranks of abnormal accruals when decile ranks of operating cash flows-to-price ratio, book-to-market ratio, and sales growth are included in the model.

[c] Zero-investment portfolio returns in year $t+1$ are calculated by subtracting the average size-adjusted return of firms in the highest abnormal accruals decile (i.e., short position) from the average size-adjusted return of firms in the lowest abnormal accruals decile (i.e., long position). The "Mean" amount is the average return to the portfolio over the sample period. The "Standard Deviation" is the standard deviation of the portfolio's returns. "Years Positive" represents the number of years (out of a possible 13) that the zero-investment portfolio produces a positive return in year $t+1$. "Excess" is calculated by taking the announcement return of each firm in the zero-investment portfolio and subtracting a random 12-day non-announcement return for the same firm in the same year. Amounts in parentheses represent averages for 100 random size-matched zero-investment portfolios (see Section II for a description of the randomization procedure).

[d] Annual returns are measured as size-adjusted returns over the 12-month period beginning three months after the fiscal year-end.

[e] Announcement returns are calculated as the 12-day size-adjusted return, consisting of the four three-day periods surrounding quarterly earnings announcements in year $t+1$.

modified Jones model and decile rank measures of abnormal accruals in the second stage return regressions. To be complete, we also apply decile rank measures to the first stage accruals expectation model.[16] Accruals and financial variables are decile ranked within each industry in each year. The resulting abnormal accrual estimate is then ranked into deciles for the second stage return regression. Rank models help to control for possible nonlinearity between accruals and financial variables and for the undue influence that extreme observations may have. As shown in Table 5, we find that results are quite similar to those reported previously. All tests continue to support the incremental existence of the abnormal accrual anomaly.

Abnormal Accruals Estimation Using Firm-Specific Models

To this point, results are presented using cross-sectional industry-specific regressions to estimate abnormal accruals. This is comparable to the majority of prior research. In this section, we consider estimation of abnormal accruals using firm-specific models, where

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[16] We also consider using full rank models (instead of deciles). Results are similar to those reported.

TABLE 5
Relation between Abnormal Accruals and Future Returns Using Rank Models to Estimate Abnormal Accruals[a]

Regressions[b]		Zero-Investment Portfolios[c]			
Coef.	Coef. with Controls	Mean	Standard Deviation	Years Positive	Excess
Panel A: Annual Returns[d]					
0.091*	0.053**	0.098* (−0.005)	0.108 (0.076)	12 (6)	
Panel B: Announcement Returns[e]					
0.034*	0.034*	0.036* (0.000)	0.032 (0.030)	13 (7)	0.027*

*,** Significant at the 0.05 and 0.10 levels, respectively, using a one-tailed t-test that the mean is greater than zero.

[a] The sample period is 1989–2001. Accruals and operating cash flows are amounts from the statement of cash flows. Abnormal accruals are estimated using the modified Jones model within each two-digit SIC code each year using decile ranks of the dependent and independent variables.

[b] "Coef." represents the average annual coefficient in a regression of future size-adjusted returns on decile ranks of abnormal accruals. Decile ranks are 0 to 9, scaled by 9. Decile ranks are multiplied by −1. "Coef. with Controls" represents the average annual coefficient in a regression of future size-adjusted returns on decile ranks of abnormal accruals when decile ranks of operating cash flows-to-price ratio, book-to-market ratio, and sales growth are included in the model.

[c] Zero-investment portfolio returns in year $t+1$ are calculated by subtracting the average size-adjusted return of firms in the highest abnormal accruals decile (i.e., short position) from the average size-adjusted return of firms in the lowest abnormal accruals decile (i.e., long position). The "Mean" amount is the average return to the portfolio over the sample period. The "Standard Deviation" is the standard deviation of the portfolio's returns. "Years Positive" represents the number of years (out of a possible 13) that the zero-investment portfolio produces a positive return in year $t+1$. "Excess" is calculated by taking the announcement return of each firm in the zero-investment portfolio and subtracting a random 12-day non-announcement return for the same firm in the same year. Amounts in parentheses represent averages for 100 random size-matched zero-investment portfolios (see Section II for a description of the randomization procedure).

[d] Annual returns are measured as size-adjusted returns over the 12-month period beginning three months after the fiscal year-end.

[e] Announcement returns are calculated as the 12-day size-adjusted return, consisting of the four three-day periods surrounding quarterly earnings announcements in year $t+1$.

firm observations over the entire sample period are employed. If the relation between financial variables and normal accruals is firm-specific, then industry-specific models lead to measurement error. However, to the extent that the relation between financial variables and normal accruals is also time-specific, firm-specific estimation will contain measurement error.

In addition, we note that firm-specific estimation is problematic for our tests of mispricing for two reasons. First, data from the statement of cash flows have been provided for a relatively small number of years. In generating abnormal accruals, we restrict each model by requiring at least three degrees of freedom.[17] The low number of time-series observations reduces the efficiency in estimating abnormal accruals. Second, and perhaps more importantly, *ex post* data must be used (i.e., data for the entire sample period are used to estimate abnormal accruals in each year). This introduces a potential look-ahead bias since data used to calculate abnormal accruals are not available to investors at the time they are setting prices. We caution the reader to keep these potential shortcomings in mind when interpreting the results.[18]

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[17] This reduces the overall sample to 16,906 firm/year observations.

As shown in Table 6, conclusions remain when using firm-specific estimation. Abnormal accruals continue to relate to future returns incremental to OCF/P, and this relation is especially apparent during earnings announcements. We also note that if a look-ahead bias exists because of using firm-specific observations over the entire sample period to estimate abnormal accruals, then abnormal returns should be more pronounced in earlier years. Closer examination of the performance of our firm-specific models (untabulated) shows that this is not the case. As an additional robustness test, we perform out-of-sample tests by using past information only. We restrict our analyses from 13 years to the last five years, requiring at least eight years of past data to estimate firm-specific parameters. Using pooled regressions over this smaller sample period (untabulated), we continue to find evidence that abnormal accruals relate to future returns incremental to OCF/P.

TABLE 6

Relation between Abnormal Accruals and Future Returns Using Firm-Specific Models to Estimate Abnormal Accruals[a]

Regressions[b]		**Zero-Investment Portfolios**[c]			
Coef.	**Coef. with Controls**	**Mean**	**Standard Deviation**	**Years Positive**	**Excess**
Panel A: Annual Returns[d]					
0.117*	0.069*	0.137* (−0.008)	0.067 (0.083)	13 (6)	
Panel B: Announcement Returns[e]					
0.032*	0.030*	0.043* (0.002)	0.026 (0.035)	13 (7)	0.044*

* Significant at the 0.05 level using a one-tailed t-test that the mean is greater than zero.

[a] The sample period is 1989–2001. Accruals and operating cash flows are amounts from the statement of cash flows. Abnormal accruals are estimated using the modified Jones model within each firm over the entire sample period.

[b] "Coef." represents the average annual coefficient in a regression of future size-adjusted returns on decile ranks of abnormal accruals. Decile ranks are 0 to 9, scaled by 9. Decile ranks are multiplied by −1. "Coef. with Controls" represents the average annual coefficient in a regression of future size-adjusted returns on decile ranks of abnormal accruals when decile ranks of operating cash flows-to-price ratio, book-to-market ratio, and sales growth are included in the model.

[c] Zero-investment portfolio returns in year $t+1$ are calculated by subtracting the average size-adjusted return of firms in the highest abnormal accruals decile (i.e., short position) from the average size-adjusted return of firms in the lowest abnormal accruals decile (i.e., long position). The "Mean" amount is the average return to the portfolio over the sample period. The "Standard Deviation" is the standard deviation of the portfolio's returns. "Years Positive" represents the number of years (out of a possible 13) that the zero-investment portfolio produces a positive return in year $t+1$. "Excess" is calculated by taking the announcement return of each firm in the zero-investment portfolio and subtracting a random 12-day non-announcement return for the same firm in the same year. Amounts in parentheses represent averages for 100 random size-matched zero-investment portfolios (see text for description of the randomization procedure).

[d] Annual returns are measured as size-adjusted returns over the 12-month period beginning three months after the fiscal year-end.

[e] Announcement returns are calculated as the 12-day size-adjusted return, consisting of the four three-day periods surrounding quarterly earnings announcements in year $t+1$.

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[18] See DeFond and Jiambalvo (1994) and Xie (2001) for examples of studies that employ both industry-specific and firm-specific estimation.

V. SUMMARY

Desai et al. (2004) provide evidence that the ability of accruals to explain future returns is subsumed by the operating cash flows-to-price ratio (OCF/P). They suggest that returns to an accrual strategy represent returns to the overall value-glamour anomaly and do not likely represent the mispricing of earnings. We extend this line of research in two ways. First, we employ multiple measures of abnormal accruals. Second, we test the relation between (abnormal) accruals and both future annual returns and future earnings announcement returns. Returns to risk-based strategies are more likely to occur smoothly over the year (Bernard et al. 1997). In contrast, returns related to the mispricing of earnings are more likely to concentrate around events related to realizations of future earnings (e.g., year-ahead quarterly earnings announcements). Thus, an analysis of announcement returns relative to annual returns may provide deeper insights into the anomalous behavior of accruals.

We employ a number of tests of the relation between abnormal accruals and future returns, including univariate and multivariate regression analyses and zero-investment portfolio analyses. We also compare the results of our zero-investment portfolio analyses to randomly generated returns for size-matched portfolios. The procedure involves randomly matching each firm in the long or short position of the abnormal accrual strategy to a firm in the same size decile in the same year. We then compare annual returns and announcement returns (e.g., means, standard deviations, and number of years that have positive returns) of the accrual strategy to those of the random size-matched strategy. To provide further evidence of anomalous behavior, we compare the 12-day announcement return to random 12-day non-announcement returns.

Consistent with results reported in Desai et al. (2004), we find that OCF/P subsumes *total* accruals in explaining future annual return. However, we note that OCF/P does not subsume *total* accruals in explaining future *announcement* returns. While future returns to the OCF/P strategy occur smoothly over the following year, returns to the total accrual strategy cluster around dates corresponding to the release of future earnings information. We find that *abnormal* accruals, our variable of primary interest, are not subsumed by OCF/P. Abnormal accruals continue to have significant explanatory power for future *annual* returns, even after controlling for OCF/P. Furthermore, we find that returns to the abnormal accrual strategy concentrate around quarterly earnings announcements. In a multiple regression of future announcement returns on abnormal accruals and control variables including OCF/P, book-to-market ratio, and sales growth, the coefficient on abnormal accruals is significant and the coefficient on OCF/P has no significance. These results make it difficult to conclude convincingly that (abnormal) accruals are purely a manifestation of the value-glamour anomaly.

Results are robust to a number of sensitivity tests. We estimate abnormal accruals with the modified Jones model (Dechow et al. 1995) using both industry-specific models each year and using firm-specific models over the sample period. We estimate abnormal accruals from both total accruals and working capital accruals. We also consider rank regressions in the estimation of abnormal accruals to control for possible nonlinear effects. In all specifications, we note that abnormal accruals have incremental explanatory power for future returns beyond OCF/P, and returns to the abnormal accrual strategy are disproportionately high during year-ahead quarterly earnings announcements.

Our study contributes to the current debate on the existence and the extent of the accrual anomaly by providing robust evidence of the incremental relation between abnormal accruals and future returns. Moreover, the methodology employed in this paper, such as analyzing earnings announcement returns and comparing accrual-based portfolio returns to randomly matched portfolio returns, should be considered by researchers in exploring further mispricing phenomena.

REFERENCES

Bernard, V., J. Thomas, and J. Whalen. 1997. Accounting-based stock price anomalies: Separating market inefficiencies from risk. *Contemporary Accounting Research* 14 (Summer): 89–136.

Dechow, P., R. Sloan, and A. Sweeney. 1995. Detecting earnings management. *The Accounting Review* 70 (April): 193–225.

———, and ———. 1997. Returns to contrarian investment strategies: Tests of naive expectations hypotheses. *Journal of Financial Economics* 43 (January): 3–27.

———, and I. Dichev. 2002. The quality of accounting and earnings: The role of accruals estimation errors. *The Accounting Review* 77 (Supplement): 35–59.

DeFond, M., J. Jiambalvo. 1994. Debt covenant violation and manipulation of accruals. *Journal of Accounting and Economics* 17 (January): 145–176.

Desai, H., S. Rajgopal, and M. Venkatachalam. 2004. Value-glamour and accruals mispricing: One anomaly or two? *The Accounting Review* 79 (April): 355–385.

Doukas, J., C. Kim, and C. Pantzalis. 2002. A test for the errors-in-expectations explanation of the value/glamour stock returns performance: Evidence from analysts' forecasts. *Journal of Finance* 57 (October): 2,143–2,165.

Fama, E., and K. French. 1992. The cross-section of expected stock returns. *Journal of Finance* 47 (June): 427–466.

Hribar, P., and D. Collins. 2002. Errors in estimating accruals: Implications for empirical research. *Journal of Accounting Research* 40 (March): 105–134.

Jones, J. 1991. Earnings management during import relief investigations. *Journal of Accounting Research* 29 (Autumn): 193–228.

Kothari, S. P., A. Leone, and C. Wasley. 2005. Performance matched discretionary accruals models. *Journal of Accounting and Economics* 39 (February): 163–197.

Kraft, A., A. Leone, and C. Wasley. 2006. An analysis of the theories and explanations offered for the mispricing of accruals and accruals components. *Journal of Accounting Research* 44 (May): 297–339.

La Porta, R., J. Lakonishok, A. Shleifer, and R. Vishny. 1997. Good news for value stocks: Further evidence on market efficiency. *Journal of Finance* 52 (June): 859–874.

Lakonishok, J., A Shleifer, and R. Vishny. 1994. Contrarian investment, extrapolation, and risk. *Journal of Finance* 49 (December): 1,541–1,578.

McNichols, M. 2002. Discussion of: The quality of accounting and earnings: The role of accruals estimation errors. *The Accounting Review* 77 (Supplement): 61–69.

Sloan, R. 1996. Do stock prices fully reflect information in accruals and cash flows about future earnings? *The Accounting Review* 71 (July): 289–315.

Xie, H. 2001. The mispricing of abnormal accruals. *The Accounting Review* 76 (July): 357–373.

Zach, T. 2003. Inside the "accruals anomaly." Working paper, Washington University.

Editorial Data

The following table contains information about turnaround time for manuscripts (including revisions) on which editorial decisions were made in the 12-month period ended December 31, 2005. Turnaround time is the number of days between the date that the manuscript was received *and the date of the editor's letter to the author(s).*

	Number of Manuscripts	Percent	Cumulative Number	Cumulative Percent
0 ≤ Days ≤ 30	84	18.06	84	18.06
31 ≤ Days ≤ 60	183	39.36	267	57.42
61 ≤ Days ≤ 90	140	30.11	407	87.53
91 ≤ Days ≤ 120	49	10.54	456	98.07
121 ≤ Days	09	1.93	465	100.00
Total	465	100.00		

The mean review time was 55.93; the median review time was 77 days.

New Submissions by calendar year (January 1 to December 31)

1992	241
1993	234
1994	231
1995	195
1996	230
1997	195
1998	196
1999	239
2000	260
2001	260*
2002	324
2003	327
2004	307
2005	389

* 2001—in addition to the 260 regular submissions, 68 MSs were submitted to the *TAR* Quality of Earnings Conference.

The acceptance rate (defined as number of MSs accepted divided by the number of new submissions) over the past 5 years is between 12–16% per year.

EDITORIAL POLICY AND STYLE INFORMATION

EDITORIAL POLICY

According to the policies set by the Publications Committee (which were endorsed by the Executive Committee and were published in the *Accounting Education News*, June 1987), *The Accounting Review* "should be viewed as the premier journal for publishing articles reporting the results of accounting research and explaining and illustrating related research methodology. The scope of acceptable articles should embrace any research methodology and any accounting-related subject, as long as the articles meet the standards established for publication in the journal ... No special sections should be necessary. The primary, but not exclusive, audience should be—as it is now—academicians, graduate students, and others interested in accounting research."

The primary criterion for publication in *The Accounting Review* is the significance of the contribution an article makes to the literature.

The efficiency and effectiveness of the editorial review process is critically dependent upon the actions of both authors submitting papers and the reviewers. Authors accept the responsibility of preparing research papers at a level suitable for evaluation by independent reviewers. Such preparation, therefore, should include subjecting the manuscript to critique by colleagues and others and revising it accordingly prior to submission. The review process is not to be used as a means of obtaining feedback at early stages of developing the research.

Reviewers and associate editors are responsible for providing critically constructive and prompt evaluations of submitted research papers based on the significance of their contribution and on the rigor of analysis and presentation. Associate editors also make editorial recommendations to the editor.

MANUSCRIPT PREPARATION AND STYLE

The Accounting Review's manuscript preparation guidelines follow *The Chicago Manual of Style* (15th ed.; University of Chicago Press). Another helpful guide to usage and style is *The Elements of Style*, by William Strunk, Jr., and E. B. White (Macmillan). Spelling follows *Webster's Collegiate Dictionary.*

FORMAT

1. All manuscripts should be printed in 12-point font on one side of $8\frac{1}{2} \times 11''$ good quality paper and be double-spaced, except for indented quotations.
2. Manuscripts should be as concise as the subject and research method permit, generally not to exceed 7,000 words.
3. Margins of at least one inch from top, bottom, and sides should facilitate editing and duplication.
4. To promote anonymous review, authors should not identify themselves directly or indirectly in their papers or in experimental test instruments included with the submission. Single authors should not use the editorial "we."
5. A cover page should show the title of the paper, the author's name, title and affiliation, email address, any acknowledgments, and a footnote indicating whether the author would be willing to share the data (see last paragraph in this statement).

Pagination: All pages, including tables, appendices and references, should be serially numbered. Major sections should be numbered in Roman numerals. Subsections should not be numbered.

Numbers: Spell out numbers from one to ten, except when used in tables and lists, and when used with mathematical, statistical, scientific, or technical units and quantities, such as distances, weights and measures. For example: *three days; 3 kilometers; 30 years.* All other numbers are expressed numerically.

Percentages and Decimal Fractions: In nontechnical copy use the word percent in the text.

Hyphens: Use a hyphen to join unit modifiers or to clarify usage. For example: *a well-presented analysis; re-form.* See *Webster's* for correct usage.

Keywords: The abstract is to be followed by four keywords that will assist in indexing the paper.

ABSTRACT/INTRODUCTION

An Abstract of about 100 words should be presented on a separate page immediately preceding the text. The Abstract should concisely inform the reader of the manuscript's topic, its methods, and its findings. Keywords and the Data Availability statements should follow the Abstract. The text of the paper should start with a section labeled "I. Introduction," which provides more details about the paper's purpose, motivation, methodology, and findings. Both the Abstract and the Introduction should be relatively nontechnical, yet clear enough for an informed reader to understand the manuscript's contribution. The manuscript's title, but neither the author's name nor other identification designations, should appear on the Abstract page.

TABLES AND FIGURES

The author should note the following general requirements:

1. Each table and figure (graphic) should appear on a separate page and should be placed at the end of the text. Each should bear an Arabic number and a complete title indicating the exact contents of the table or figure. Tables and figures should define each variable. The titles and definitions should be sufficiently detailed to enable the reader to interpret the tables and figures without reference to the text.
2. A reference to each graphic should be made in the text.
3. The author should indicate by marginal notation where each graphic should be inserted in the text.
4. Graphics should be reasonably interpreted without reference to the text.
5. Source lines and notes should be included as necessary.

Equations: Equations should be numbered in parentheses flush with the right-hand margin.

DOCUMENTATION

Citations: Work cited should use the "author-date system" keyed to a list of works in the reference list (see below). Authors should make an effort to include the relevant page numbers in the cited works.

1. In the text, works are cited as follows: authors' last name and date, without comma, in parentheses: for example, (Jones 1987); with two authors: (Jones and Freeman 1973); with more than two: (Jones et al. 1985); with more than one source cited together (Jones 1987; Freeman 1986); with two or more works by one author: (Jones 1985, 1987).
2. Unless confusion would result, do not use "p." or "pp." before page numbers: for example, (Jones 1987, 115).
3. When the reference list contains more than one work of an author published in the same year, the suffix a, b, etc. follows the date in the text citation: for example, (Jones 1987a) or (Jones 1987a; Freeman 1985b).
4. If an author's name is mentioned in the text, it need not be repeated in the citation; for example, "Jones (1987, 115) says ..."
5. Citations to institutional works should use acronyms or short titles where practicable; for example, (AAA ASOBAT 1966); (AICPA Cohen Commission Report 1977). Where brief, the full title of an institutional work might be shown in a citation: for example, (ICAEW *The Corporate Report* 1975).
6. If the manuscript refers to statutes, legal treatises or court cases, citations acceptable in law reviews should be used.

Reference List: Every manuscript must include a list of references containing only those works cited. Each entry should contain all data necessary for unambiguous identification. With the author-date system, use the following format recommended by *The Chicago Manual*:

1. Arrange citations in alphabetical order according to surname of the first author or the name of the institution responsible for the citation.
2. Use author's initials instead of proper names.
3. Date of publication should be placed immediately after author's name.
4. Titles of journals should not be abbreviated.
5. Multiple works by the same author(s) in the same year are distinguished by letters after the date.
6. Inclusive page numbers are treated as recommended in *Chicago Manual* section 17.168.B.

Sample entries are as follows:

American Accounting Association, Committee on Concepts and Standards for External Financial Reports. 1977. *Statement on Accounting Theory and Theory Acceptance.* Sarasota, FL: AAA.

Demski, J. S., and D. E. M. Sappington. 1989. Hierarchical structure and responsibility accounting. *Journal of Accounting Research* 27 (Spring): 40–58.
Dye, R., B. Balachandran, and R. Magee. 1989. Contingent fees for audit firms. Working paper, Northwestern University, Evanston, IL.
Fabozzi, F., and I. Pollack, eds. 1987. *The Handbook of Fixed Income Securities.* 2nd edition. Homewood, IL: Dow Jones-Irwin.
Kahneman, D., P. Slovic, and A. Tversky, eds. 1982. *Judgment Under Uncertainty: Heuristics and Biases.* Cambridge, U.K.: Cambridge University Press.
Porcano, T. M. 1984a. Distributive justice and tax policy. *The Accounting Review* 59 (October): 619–636.
———. 1984b. The perceived effects of tax policy on corporate investment intentions. *The Journal of the American Taxation Association* 6 (Fall): 7–19.
Shaw, W. H. 1985. Empirical evidence on the market impact of the safe harbor leasing law. Ph.D. dissertation, The University of Texas at Austin.
Sherman, T. M., ed. 1984. *Conceptual Framework for Financial Accounting.* Cambridge, MA: Harvard Business School.

Footnotes: Footnotes are not used for documentation. Textual footnotes should be used only for extensions and useful excursions of information that if included in the body of the text might disrupt its continuity. Footnotes should be consecutively numbered throughout the manuscript with superscript Arabic numerals. Footnote text should be doubled-spaced and placed at the end of the article.

SUBMISSION OF MANUSCRIPTS

Authors should note the following guidelines for submitting manuscripts:

1. Manuscripts currently under consideration by another journal or publisher should not be submitted. The author must state that the work is not submitted or published elsewhere.
2. For manuscripts reporting on field surveys or experiments: If the additional documentation (e.g., questionnaire, case, interview schedule) is sent as a separate file, then all information that might identify the author(s) must be deleted from the instrument.
3. MS should be submitted via email as Microsoft® Word or PDF files to the *TAR* office at TARUAZ@email.arizona.edu. A submission letter, a separate cover page, and the MS containing no author identification should be emailed as separate files. The submission fee should be paid online (VISA or MasterCard only) at: https://aaahq.org/AAAforms/journals/tarsubmit.cfm. If unable to pay online, a check in U.S. funds, made payable to the American Accounting Association, may be sent to Professor Dan Dhaliwal, Senior Editor, *The Accounting Review*, The University of Arizona, Department of Accounting, McClelland Hall 301, 1130 E. Helen Street, Tucson, AZ 85721-0108. The submission fee in U.S. funds is $125.00 for members and $200.00 for nonmembers of the AAA. Effective January 1990, the submission fee is nonrefundable.
4. The author should retain a copy of the paper.
5. Revisions must be submitted within 12 months from request, otherwise they will be considered new submissions.

COMMENTS

Comments on articles previously published in *The Accounting Review* will be reviewed (anonymously) by two reviewers in sequence. The first reviewer will be the author of the original article being subjected to critique. If substance permits, a suitably revised comment will be sent to a second reviewer to determine its publishability in *The Accounting Review*. If a comment is accepted for publication, the original author will be invited to reply. All other editorial requirements, as enumerated above, also apply to proposed comments.

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POLICY ON DATA AVAILABILITY

The following policy has been adopted by the Executive Committee in its April 1989 meeting. "An objective of (*The Accounting Review, Accounting Horizons, Issues in Accounting Education*) is to provide the widest possible dissemination of knowledge based on systematic scholarly inquiries into accounting as a field of professional research, and educational activity. As part of this process, authors are encouraged to make their data available for use by others in extending or replicating results reported in their articles. Authors of articles which report data dependent results should footnote the status of data availability and, when pertinent, this should be accompanied by information on how the data may be obtained."

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OCTOBER 2006 PLACEMENT ADS

*The deadline for information to be included in the "Placement Information" section is two months prior to the date of publication. The **Review** is published in January, March, May, July, and October. Placement advertising and replies should be sent to the American Accounting Association, 5717 Bessie Drive, Sarasota, FL 34233-2399, ATTENTION: ADVERTISING DEPT. All replies to box numbers should be on separate sheets to facilitate forwarding.*

501 UNIVERSITY OF SOUTHERN INDIANA, College of Business invites applications for a tenure-track position in Accounting beginning August 2007. Candidates seeking the rank of Assistant Professor must have a Ph.D. or D.B.A. in Accounting (or a related discipline). ABDs will be considered at the rank of Instructor until completion of the degree. Preference will be given to candidates with a primary specialization in Financial Accounting (although other areas will be considered). Responsibilities include teaching in multiple accounting areas. The search will continue until the position is filled, but preference will be given to applications and nominations received by November 15, 2006. Apply (electronically or by mail) with a letter of application, *curriculum vitae*, teaching philosophy statement, and evidence of teaching effectiveness and scholarship to: Linda Dillbeck, College of Business, University of Southern Indiana, 8600 University Blvd., Evansville, IN 47712; Phone: (812) 464-1926; Email: ldillbec@usi.edu. Information about USI and the College of Business can be found at: http://www.usi.edu and http://business.usi.edu. The University of Southern Indiana is an Equal Opportunity/Affirmative Action Employer.

502 MIDWESTERN STATE UNIVERSITY, Dillard College of Business Administration invites applications for a tenure-track position in Accounting, commencing Fall 2007. A Ph.D. in accounting from an accredited university, a passion for excellence in teaching, and a commitment to research consistent with AACSB standards are required. Qualified candidates will be considered for appointment to the Madera Professorship in Accounting. Preference will be given to candidates with the ability to teach auditing and related courses. ABDs will be considered. Rank is open and salary is competitive. To be considered, please send a letter of application, current *curriculum vitae*, unofficial transcripts, and list of three references to: Dr. Ralph Fritzsch, Chair, Department of Accounting, Dillard College of Business Administration, Midwestern State University, 3410 Taft, Wichita Falls, TX 76308; Email: ralph.fritzsch@mwsu.edu. Website: http://business.mwsu.edu/accounting/index.asp.

503 AUGUSTANA COLLEGE, James Galley Chair in Accounting invites applications for a tenure-track position in Accounting at the rank of Assistant or Associate Professor beginning the 2007–2008 academic year. The ideal candidate will hold a Ph.D. in accounting, hold a CPA or CMA certificate, and have two or more years of practical experience in corporate or public accounting. ABDs nearing completion and others with extraordinary qualifications are also encouraged to apply. Candidates should be able to demonstrate the ability to bring excellence and innovation into the classroom, maintain a continuing record of scholarship, and possess attributes leading to successful professional service. The position includes teaching six sections in the financial accounting curriculum, participating in department responsibilities, and performing the duties as the Galley Chair. Those duties include support of the traditional accounting curriculum while strengthening the connection of the department to the liberal arts areas on the campus. Augustana College is a more selective, four-year, liberal arts institution of 2,400 students, most of whom live in residence halls on a wooded 115-acre campus. Details about Augustana, our expectation of the faculty, the selection process, and the Quad Cities are all available at the Faculty Search website; http://www.augustana.edu/academics/faculty_search/. Send a letter of application, *curriculum vitae*, graduate transcripts, statement of teaching philosophy, and three letters of recommendation to: Accounting Search Committee, Jeff Abernathy, Dean of the College, Augustana College, 639 38th Street, Rock Island, IL 61201. Questions may be directed to abernathy@augustana.edu. Review of applications will begin June 15, 2006. Augustana College is an Equal Opportunity Employer and actively encourages applications from women and persons of diverse ethnic backgrounds. We do not discriminate based on age, race, color, ethnic origin, gender, sexual orientation, disability, or creed.

504 TEXAS A&M UNIVERSITY–CORPUS CHRISTI, College of Business, invites applications for a Visiting Instructor of Accounting. An earned master's degree in business (accounting master's is preferred) with at least 18 graduate credit hours in accounting is required. Certification and professional experience in accounting are preferred. Appointment is for the 2006–07 academic year commencing Fall 2006. Salary is competitive. Applicant review will continue until position is filled. Send applications with a list of three references to: Dr. Donald R. Deis, Chair of Search Committee, Texas A&M University–Corpus Christi, College of Business, 6300 Ocean Drive Unit #5808, Corpus Christi, TX 78412-5808; Email: Donald.Deis@tamucc.edu. College of Business undergraduate and graduate business programs are accredited by AACSB International. Texas A&M University–Corpus Christi is an Equal Opportunity/Affirmative Action Employer. Additional information about the College and its programs is available at: http://www.cob.tamucc.edu.

505 Assistant Professor with Accounting Doctorate, CPA, and CMA seeks tenure—track position in the Southeast, Texas, New Mexico, or Missouri. AACSB-qualified with recent publications. Substantial experience teaching principles, managerial, cost and intermediate. Has also taught other advanced courses. Available Fall 2006. Please reply to: acctg_prof@yahoo.com.

506 UNIVERSITY OF TENNESSEE, KNOXVILLE, is accepting applications for a full-time, tenure-track Financial Accounting position (Assistant or Associate Professor) to begin August 1, 2007. Candidates should have an interest in teaching financial and international accounting and in producing capital markets research. Accordingly, candidates should hold a doctoral degree with a specialization in the financial accounting area, or expect to complete all degree requirements by the starting date. Please send applications to: Kenneth E. Anderson, Search Committee Chair, Department of Accounting and Information Management, CBA, University of Tennessee, Knoxville, TN 37996-0560; Email: kea@utk.edu UT Knoxville is an EEO/AA/Title VI/Title IX/Section 504/ADA/ADEA institution in the provision of its education and employment programs and services.

507 VIRGINIA POLYTECHNIC INSTITUTE & STATE UNIVERSITY, Department of Accounting and Information Systems seeks to fill one or more tenure-track positions to begin August 10, 2007. All ranks will be considered. Our primary need is in Financial Accounting, but exceptionally qualified applicants specializing in Managerial Accounting, Auditing, or Taxation will be considered. Successful candidates will have a Ph.D. in accounting from an AACSB-accredited school, a program of scholarly research consistent with the rank applied for, and instructional capabilities at both the undergraduate and graduate levels. Information about the Department and its programs is available at: http://www.acis.cob.vt.edu. Interested candidates *must* apply online at: http://www.jobs.vt.edu (posting number 060485) and include with the electronic application a cover letter, resume, and the names of three references. Review of applications will begin July 1, 2006. General inquiries should be directed to: Professor Bryan Cloyd, Search Committee Chair, 3007 Pamplin Hall (0101), Blacksburg, VA 24061. Virginia Tech is an Equal Opportunity/Affirmative Action Employer.

508 WASHBURN UNIVERSITY, AACSB International-Accredited School of Business invites applications for Assistant or Associate Professor of Accounting to begin Fall 2007. Preference given to candidates with Ph.D. completed; ABDs considered if substantial progress made toward completion of dissertation. Undergraduate and M.B.A. teaching responsibilities may include financial, managerial and cost, or auditing. Ideal candidate will have demonstrated excellence in teaching, commitment to scholarship, and service commensurate with AACSB International accreditation. Submit application letter, current *vita* with three references, transcripts of all graduate work, summary of teaching performance, and copies of recent scholarly work to: Chair, Accounting Position Search Committee, School of Business, Washburn University, 1700 SW College Avenue, Topeka, KS 66621. Competitive salary dependent on qualifications and experience. Application review begins July 15, 2006 and continues until position filled. Washburn University is an Equal Opportunity Employer. Website: http://www.washburn.edu/business.

509 LOUISIANA STATE UNIVERSITY, Department of Accounting in the E. J. Ourso College of Business, invites applications for one or more anticipated faculty positions at the rank of Associate or Full Professor to begin August 2007. Applications will be considered in all areas of specialty; however, priority will be given to candidates with a specialty in financial accounting. Required Qualifications: Earned doctorate in accounting; established record of high-quality academic research; continuing research agenda; interest in guiding the future of the Ph.D. program in accounting. Additional Qualification Desired: CPA. Responsibilities: teach undergraduate and/or graduate courses; be actively involved in academic research; provide service to the profession and the University. It is anticipated that this position will assume responsibility for the doctoral program. An offer of employment is contingent on a satisfactory pre-employment background check. Applications will be accepted until April 16, 2007, or until candidate is selected. Please send a letter of application, *curriculum vitae*, email address, and three references to: Dr. Joey Legoria, Department of Accounting, 3101 CEBA Building, Louisiana State University, Ref: Log: #0659, Baton Rouge, LA 70803. Phone: (225) 578-6225, Email: jlegoria@lsu.edu. Website: http://www.bus.lsu.edu/academics/accounting/. LSU is an Equal Opportunity/Equal Access Employer.

510 WICHITA STATE UNIVERSITY, invites applications for three tenure-eligible Assistant or Associate Professor positions. Required qualifications: Ph.D./D.B.A. in Accounting (ABD considered) for Assistant; Ph.D./D.B.A. in Accounting and significant research record for Associate; teaching and research interests in accounting; commitment to quality teaching and research; commitment to interact with the professional community; and commitment to diversity. Desirable qualifications: professional accounting experience; professional certification; teaching and research interests in any area of accounting, consistent with the needs of the School of Accountancy. This nine-month appointment begins August 2007. For full consideration, submit applications by October 27, 2006. The position remains open until filled. Salary is competitive (position is contingent upon funding). Send a *curriculum vitae* and names of three references to: Jeff Bryant, School of Accountancy, Wichita State University, 1845 Fairmount, Wichita, KS 67260-0087; Phone: 316-978-6261; Email: jeffrey.bryant@wichita.edu. Wichita State University is an Equal Opportunity/Affirmative Action Employer.

511 UNIVERSITY OF MINNESOTA, DULUTH, invites applications for a full-time, nine-month, tenure-track faculty position in the Department of Accounting, University of Minnesota, Duluth beginning Fall 2007. The individual hired will be responsible for teaching, research, and service. The individual will teach and develop courses in financial/intermediate accounting, cost accounting and/or auditing. See http://www.d.umn.edu/lsbe/faculty/openpositions.php for a copy of the position announcement. Further information may be obtained from: Alan C. Roline, Chair, 125 SBE, UMD, 412 Library Drive, Duluth, MN 55812-3029; Phone: (218) 726-8550; Email: aroline@d.umn.edu; fax 218-726-8510. The University of Minnesota is an Equal Opportunity Educator and Employer.

512 BALL STATE UNIVERSITY, Department of Accounting invites applications for a tenure-track position available August 2007. Responsibilities: teaching courses in accounting information systems at the undergraduate and graduate levels. Minimum qualifications: ABD toward doctorate in accounting or other business related field, completion of degree by November 1, 2008; evidence of scholarly research; teaching experience in accounting information systems or other business related field. Preferred qualifications: Ph.D. or D.B.A. in accounting or other business-related field; full-time teaching experience at the university level; significant record of publications in refereed journals in the accounting area and/or evidence of other scholarly contributions. Send letter of application, *curriculum vitae*, official transcripts, and the names and contact information for three references to: Dr. Lucinda L. Van Alst, Chairperson, Department of Accounting, Ball State University, Muncie, IN 47306. Review of applications will begin immediately and will continue until the position is filled. Ball State University is an Equal Opportunity/Affirmative Action Employer and is strongly and actively committed to diversity within its community. Website: http://www.bsu.edu.

513 THE UNIVERSITY OF TEXAS AT AUSTIN, Department of Accounting is seeking applicants for as tenure-track position to begin September 1, 2007. Appointments will be considered at all academic ranks, including endowed chair and professorship appointments, in all areas of accounting research. Entry-level candidates should have (or be a candidate for) a Ph.D. or D.B.A. with demonstrated potential for excellence in research and teaching. Advanced position candidates must have commensurately outstanding records and potential. Candidates are expected to pursue an active research program, perform graduate and undergraduate teaching, and supervise graduate students. Excellent English communication skills are required. Submit *vita*, transcripts, references, research proposals and/or samples or research to: Dr. Ross Jennings, Chairman, The University of Texas at Austin, McCombs School of Business, Department of Accounting, 1 University Station-B6400, Austin, TX 78712-0211. To be certain of consideration, apply by January 15, 2007. The University of Texas is an Equal Opportunity/Affirmative Action Employer.

514 TSINGHUA UNIVERSITY, School of Economics and Management, invites applications in all areas of accounting to apply for the openings in the Department of Accounting beginning June 2006. All ranks will be considered. Applicants should hold (or expect to hold) a Ph.D. in accounting. We stress high research potential or achievement. Candidates must be committed to effective teaching, scholarly and professional research, and university and community service. Rank and salary will be commensurate with qualifications and experience. Applicants should send a letter of application, current *curriculum vitae*, a sample of written work and three letters of reference to: Professor Yingyi Qian, First Associate Dean at: FacultyRecruiter@em.tsinghua.edu.cn. The School offers undergraduate, Master's (including M.B.A.) and Ph.D. programs. It provides an excellent academic environment, which is significantly enhanced recently by the arrival of Special-Term Professors who have established academic records in overseas research universities. For more information, please see: http://www.em.tsinghua.edu.cn.

515 UNIVERSITY OF HOUSTON–CLEAR LAKE, School of Business, is seeking three faculty members at the rank of Assistant Professor of Accounting beginning August 2007. A doctorate in Accounting or an ABD nearing completion from an AACSB-accredited school is required. Preference will be given to candidates who are conducting discipline-based research. Applications are accepted only online at: https://jobs.uhcl.edu. To apply, please complete the online application and attach a letter of interest and *vita* online. Please send your three letters of recommendation and transcript to: Barry Marks, Chair of Accounting Search Committee, University of Houston–Clear Lake, Box 248, 2700 Bay Area Boulevard, Houston, TX 77058-1098. The University reserves the right to extend the search or not to fill the positions. The candidates must be eligible to work in the United States. UHCL is an Affirmative Action/Equal Opportunity Employer supporting workplace diversity.

516 NORTH CAROLINA STATE UNIVERSITY, Department of Accounting invites applicants for a tenure-track position for Fall 2007. A specialization in Accounting Information Systems is preferred, but Audit, Financial or Managerial areas will also be considered. A doctorate (or short-term ABD status) in Accounting from an AACSB-accredited university is required. The academic rank is open and the salary competitive. Candidates should demonstrate a strong commitment to research and teaching. NC State is a land-grant institution located in one of the ten best cities to live and work in, according to Cityrating.com. The College of Management boasts 2,800 students and offers an undergraduate and graduate degree in Accounting. Please send letter of interest and *curriculum vita* directly to: Recruiting Coordinator, Department of Accounting, North Carolina State University, Campus Box 8113, Raleigh, NC, 27695-8113. Inquires and ADA accommodations can be made via email to: acctrecruiting@ncsu.edu. AA/EOE. NC State welcomes all persons without regard to sexual orientation.

517 KOREA UNIVERSITY, Business School in Seoul, Korea invites applicants for both tenure-track and non-tenure-track faculty positions at all levels starting Spring or Fall 2007. Candidates are expected to teach two courses each semester at the graduate and/or undergraduate levels. The salary and benefits are competitive. The non-tenure-track position is for two years or longer. The non-tenure-track position is likely best suited for faculty members who plan to take a leave of absence from their institutions and wish to gain insights into the Korean business environment as well as the Seoul metropolitan environments. On campus housing can be provided. Please send letter of application, current *vitae*, and the names of two references to: Professor Gil S. Bae, Coordinator in Accounting Area, School of Business Administration, Korea University, Sungbukku, Seoul, Korea 136-701; Email: gilbae@korea.ac.kr; Phone: +82-2-3290-1951. Please visit: http://biz.korea.ac.kr for more information on the School of Business Administration, Korea University.

518 UNIVERSITY OF HAWAII AT HILO, College of Business and Economics, is recruiting a tenure-track Assistant Professor of Accounting, Position #82775, general funds, full-time; tenure-track; nine-month appointment to begin approximately August 2007, pending position clearance and available funding. The University reserves the right to hire at another rank when the selected candidate is qualified for that rank. Minimum qualifications: Ph.D. in Accounting or closely related discipline from AACSB International-accredited college/university, evidence of ability to teach effectively at undergraduate level, evidence of commitment to ongoing scholarship in accounting; ability to communicate clearly are required. All areas will be considered. Salary is competitive. For full position description, application information and deadline, please go to: http://www.uhh.hawaii.edu/uhh/hr/jobs.php. Review of applications will continue until the position is filled. UH Hilo is an EEO/AA Employer D/M/V/W.

519 UNIVERSITY OF NOTRE DAME, Department of Accountancy, invites applications for position openings at the Assistant and Associate Professor levels. All research areas will be considered. Applicants for Assistant Professor positions should expect to complete doctoral degree requirements before the start of the contract period. All candidates should have a strong commitment to scholarly research and excellence in teaching. Salary is very competitive and research support is excellent. Please send *curriculum vitae* and an example of scholarly work to Professor Tom Schaefer, Chair, Department of Accountancy, Room 102D, Mendoza College of Business Administration, University of Notre Dame, Notre Dame, IN 46556-5646. The University of Notre Dame is an Equal Opportunity/Affirmative Action Employer.

520 UNIVERSITY OF ILLINOIS AT SPRINGFIELD, invites applications for tenure-track faculty positions in Accountancy at Assistant or Associate Professor rank. Salary and rank are AACSB competitive and are commensurate with teaching, research, and service experience. Ph.D. required, ABD considered. Professional certification such as CPA, CMA, CIA, CFM and professional experience required. Financial Accounting, AIS, or Governmental Accounting are preferred teaching areas. Review of applications is ongoing and will continue until the position is filled. Submit your letter of application, *vita*, and the names of three references to: Dr. Ronald McNeil, Dean, College of Business and Management, University of Illinois at Springfield, One University Plaza, Springfield IL 62703-5407. Inquiries should be addressed to: Leonard Branson, Search Committee Chair: Department of Accountancy; Phone: (217) 206-6299; Email: branson.leonard@uis.edu. UIS is an Affirmative Action/Equal Opportunity Employer with a strong institutional commitment to recruitment and retention of a diverse and inclusive campus community. Women, minorities, veterans, and persons with disabilities are encouraged to apply. For additional information about the university or its policies, please refer to our website at: http://www.uis.edu.

521 MONTANA STATE UNIVERSITY, College of Business seeks applicants to fill one tenure-track position in Accounting for Fall 2007. Montana State University has an enrollment of approximately 12,000 students and is located in Bozeman, an attractive, culturally alive small city, 90 miles north of Yellowstone National Park. Fully accredited by the AACSB, the College of Business offers a Bachelor of Science in Business with options in Accounting, Finance, Management, and Marketing, and a Master of Professional Accountancy. Colleagues in all areas of accounting will be considered; however, our greatest teaching needs are in the area of accounting theory/research. In addition, research activities and service consistent with the College's mission are expected. The position requires a Ph.D. or D.B.A. in Accounting from an AACSB-accredited university. Preference will be given to those who have completed their doctoral degrees or will complete them prior to beginning work at Montana State University. For more information and complete job descriptions with application requirements and deadlines, please visit: http://www.montana.edu/msuinfo/jobs/faculty/ or contact: Anne L. Christensen at: annec@montana.edu or 406-994-2043. ADA/AA/EO/Veteran's Preference.

522 UNIVERSITY OF NEW ORLEANS, invites applications for an anticipated Assistant Professor position for Spring 2007 or Fall 2007. Preference will be given to those with teaching interests in some combination of taxation, accounting systems, auditing, financial, and managerial accounting. Candidates must have the Ph.D. with an accounting major (by May 2007 for Spring hires and December 2007 for Fall hires). Salary is negotiable and competitive. The Department offers the undergraduate accounting degree (about 500 majors) and two Master's degrees (including a Tax Option). The Department has separate AACSB accounting accreditation at both the graduate and undergraduate levels. Candidates should send a resume to: Professor Philip J. Harmelink, Chairman, Department of Accounting, College of Business Administration, University of New Orleans, New Orleans, LA 70148; Email: pharmeli@uno.edu. The University of New Orleans is a member of the Louisiana State University System and an Affirmative Action/Equal Opportunity Employer.

523 FAIRFIELD UNIVERSITY, Charles F. Dolan School of Business (AACSB-accredited) of invites applications for a tenure-track position as Assistant/Associate Professor of Accounting for Fall 2007. Preference will be given to those in Tax/Managerial/AIS or those with an ability to teach in multiple areas. A CPA is required. A doctoral degree is strongly preferred; candidates with an M.B.A. or M.S.T. and significant professional experience and suitable research may be considered. Candidates will be expected to demonstrate a commitment to teaching excellence, quality research, and service. Fairfield University is a comprehensive Jesuit University located on a 200-acre campus in Connecticut, one hour from New York City. Please send a letter of application, *vita*, transcripts, and contact information for three references, in writing or electronically, to: Dr. Dawn Massey Chair, Accounting Department, Fairfield University, Fairfield, CT 06824-5195; Email: dmassey@mail.fairfield.edu. Fairfield University is an Affirmative Action/Equal Opportunity Employer. Women/minorities/persons with disabilities are strongly encouraged to apply.

524 MIAMI UNIVERSITY (OHIO), Department of Accountancy invites applicants for tenure-track positions at the assistant or associate professor level beginning August 2007. Applicants interested in any area of accounting will be considered. In addition, Miami is recruiting to fill the Jennifer Petters Chair in Asian Business. Miami's accounting programs are AACSB-accredited and nationally ranked. Applicants must have a doctorate or dissertation in process with expectation of completion of degree by December 31, 2007. Professional certification and experience are desirable. Faculty members are expected to provide both high-quality teaching and research. Salary and benefits are competitive. Please submit your resume and letters from three references to: Dr. Marc A. Rubin, Chairman, Department of Accountancy, Farmer School of Business, Miami University, Oxford, Ohio 45056; or Email: rubinma@muohio.edu. Application screening for interviews will begin immediately. Additional information is available at: http://www.sba.muohio.edu. Miami University is an Affirmative Action/Equal Opportunity Employer.

525 UNIVERSITY OF GEORGIA, J. M. Tull School of Accounting is seeking applications for a tenure-track position in Accounting at the Assistant or Associate Professor level. Preference will be given to applicants with teaching interests in financial accounting. The anticipated date of appointment is Fall 2007, subject to budgetary and academic approvals. Candidates for this position should have either a Ph.D. in Accounting or be well along in the dissertation stage. Applicants must have a commitment to excellence in teaching and scholarship. Salary and benefits are competitive. Applications received before November 1, 2006, are assured of consideration. Please submit application materials (current *curriculum vitae* and dissertation proposal or example of scholarly work) to: Professor Ben Ayers, J. M. Tull School of Accounting, Terry College of Business, 255 Brooks Hall, University of Georgia, Athens, GA 30602. The University of Georgia is an Equal Opportunity/Affirmative Action Institution.

526 NORTHEASTERN UNIVERSITY, College of Business Administration is seeking applications for the Joseph M. Golemme Research Professorship in Accounting. The Golemme Professorship was established in honor of Professor Joseph M. Golemme, the founder and guiding force behind Northeastern's nationally recognized Graduate School of Professional Accounting. The Professorship is a tenured faculty position, renewable every five years, conditional on performance in accordance with expectations. A qualified applicant—at the Full Professor level—is expected to be actively engaged in research and publication and to provide research leadership. In addition, the successful applicant is expected to engage in professional and academic activities in order to enhance Northeastern's visibility at the national/international level. Further, the applicant is also expected to teach in the Graduate School of Professional Accounting and in other degree programs in the College; teaching will be at a reduced level. A Ph.D./D.B.A. and a significant research and publication record in top-quality journals are required. Please send applications with a current resume to: Marjorie Platt, Accounting Group Coordinator, College of Business Administration, Northeastern University, 404 Hayden Hall, 360 Huntington Avenue, Boston, MA 02115; or Email: m.platt@neu.edu. Northeastern University is an Equal Opportunity/Affirmative Action, Title IX University.

527 NORTHEASTERN UNIVERSITY, has openings for two tenure-track positions effective September 2007 at the Assistant or Associate Professor rank in Accounting. We seek individuals holding a Ph.D. in Accounting with a commitment to excellence in teaching, research, and professional service. For the Associate Professor rank, candidates with an established publication record, teaching experience in financial, managerial or auditing, and some business background are preferred. The position offers competitive compensation and teaching responsibilities that enable a notable level of high-impact research. Rank and salary are dependent on qualifications. Support is available in terms of summer funding, graduate assistants, and travel. Accounting courses are offered at both the undergraduate and graduate levels, with graduate teaching in the M.B.A., M.S./M.B.A., M.S.A., and M.S. Tax programs. Send a letter of interest, a *curriculum vitae* including three references, and two samples of scholarly work prior to: Dr. Marjorie Platt, Accounting Group, 404 Hayden Hall, Northeastern University, Boston, MA 02115; Phone: (617) 373-4647; or electronically to: Email: m.platt@neu.edu. Northeastern University is an Affirmative Action/Equal Opportunity/Title IX Employer and particularly welcomes applications from minorities, women, and persons with disabilities.

528 SIMON FRASER UNIVERSITY, invites applications for two tenure-track appointments: Managerial Accounting and Financial Accounting. Ph.D. required; accounting designation preferred. Rank: open to the level of Associate Professor. Salary: competitive within Canada. Successful candidates require strong research record and teaching abilities. All qualified candidates are encouraged to apply; positions are subject to budgetary approval. Under the authority of the University Act personal information that is required by the University for academic appointment competitions will be collected. For further details see: http://www.sfu.ca/vpacademic/Faculty_Openings/Collection_Notice.html. Start Date: September 2007. Send *curriculum vitae* (electronic format, pdf) to: busappts@sfu.ca, and three signed original letters of reference to: Dr. Irene M. Gordon, Chair, Appointments Committee Faculty of Business Administration Simon Fraser University Burnaby, B.C. V5A 1S6 CANADA. Search closes January 31, 2007. Canadians and permanent residents given priority. SFU is committed to the principle of equity in employment and offers equal employment opportunities to qualified applicants.

529 UNIVERSITY OF PENNSYLVANIA, invites applications for positions available July 2007 in Financial, Managerial, and Tax Accounting at all academic ranks. A Ph.D. or expected completion by June 30, 2008 is required. Demonstrated ability or potential for quality teaching and research are important. Interested persons should send *vitae* including education, professional experience, publications, personal data, and, if applicable, a dissertation proposal and abstract to: Professor John Core, Accounting Department, University of Pennsylvania, Suite 1300 Steinberg-Dietrich Hall, Philadelphia, PA 19104-6365. The deadline for applications is January 15, 2007. The University of Pennsylvania values diversity and seeks talented students, faculty, and staff from diverse backgrounds. The University of Pennsylvania is an Equal Opportunity/Affirmative Action Employer. Women, minority candidates, veterans, and individuals with disabilities are strongly encouraged to apply.

530 DREXEL UNIVERSITY, LeBow College of Business, Department of Accounting and Taxation invites applications for Full and Associate Professor. Qualified candidates should be recognized scholars and have strong interest in supervising doctoral students and teaching at the doctoral level. Located in Philadelphia's University City area, Drexel's LeBow College of Business is AACSB-accredited at the graduate and undergraduate levels and offers doctorate, M.B.A., Executive M.B.A., and undergraduate degrees. LeBow has been recognized by *BusinessWeek*, *Financial Times*, *U.S. News & World Report*, and *Entrepreneur* for superior programs and the College is in a period of exponential growth in quality, achievement, and reputation. Send application materials to: David R. Campbell, Head, Department of Accounting and Taxation, LeBow College of Business, Drexel University, 3141 Chestnut Street, Philadelphia, PA 19104. Send electronic submissions to: davidcampbell@drexel.edu. Drexel University is an Affirmative Action/Equal Opportunity Employer.

531 UNIVERSITY OF WISCONSIN–WHITEWATER, Accounting Department anticipates an opening for a full-time, tenure-track position for Fall 2007. The minimum qualification is an earned terminal degree (Ph.D. or D.B.A.) in accounting or a related area. Preferred qualifications include the ability to teach Financial or Managerial/Cost Accounting, an established research record, and a professional certification (e.g., CPA or CMA). Responsibilities include teaching undergraduate and/or graduate accounting courses, along with research/scholarly activities, advising, and service. Salary is competitive and negotiable. Review of applications will begin October 9, 2006, and continue until the position is filled. All interested persons should apply in writing to: Dr. Alka Arora, Department of Accounting, C5023, University of Wisconsin–Whitewater, 800 West Main Street, Whitewater, WI 53190-1790. The University of Wisconsin–Whitewater is an Affirmative Action/Equal Opportunity Employer.

532 FLORIDA INTERNATIONAL UNIVERSITY, School of Accounting is seeking four faculty members for tenure-track positions beginning Spring 2007 or earlier. Appointments at the Assistant and Associate Professor ranks will be considered. Preferred emphasis in financial accounting, managerial accounting, and taxation. Contact Professor Christos Koulamas at (305) 348-2861. Applicant review date is October 1, 2006 or earlier. FIU is an Equal Opportunity/Equal Access/Affirmative Action Employer.

533 SEOUL NATIONAL UNIVERSITY, College of Business Administration, invites applications for a non-tenured accounting faculty position at all levels starting Fall 2006. All areas of accounting will be considered. The contract is for one year and renewable. The successful candidate is expected to teach two courses each semester at graduate and/or undergraduate levels. The position is likely best suited for a faculty member who plans to take a sabbatical leave from his/her institution and wishes to gain insights on Korean business practices and academic environments. A faculty apartment with nominal rents will be available on the campus. Please send an application, current *curriculum vitae*, and three references to Professor Tae-Sik Ahn, Associate Dean for Academic Affairs, College of Business Administration, Seoul National University, Kwanak-ku, Seoul, Korea 151-742; Phone: (82-2) 880-6901; Email: ahnts@snu.ac.kr. Please visit http://cba.snu.ac.kr for more information on College of Business Administration, Seoul National University.

534 MORAVIAN COLLEGE, has a tenure-track position in Accounting beginning Fall Term 2007. Possible teaching assignments are courses in: Cost Accounting, Accounting Information Systems, Intermediate Accounting, Federal Income Tax, Business Valuation, Not-for-Profit Accounting, Advanced Accounting, and Auditing. Experience in manufacturing, technology, and ERP software preferred. Candidates must be willing to participate fully in a liberal arts environment and be committed to teaching, scholarship, and service. Required qualifications: M.B.A. and one professional certification (CPA, CMA, CIA, CFA, CFE, or CISA). Doctorate in accounting or related field preferred. Candidates should identify their strengths or experiences in this area. Applications due: November 1, 2006. Send letter, *curriculum vitae*, evidence of teaching ability, and three letters of reference to: John D. Rossi, III, Search Committee Chair, Department of Economics and Business, Moravian College, 1200 Main Street, Bethlehem, PA 18018-6650. No email submissions. The College is committed to increasing the diversity of its community and curriculum.

535 PEPPERDINE UNIVERSITY, Graziadio School of Business and Management seeks an experienced Assistant or Associate Professor in Accounting starting Fall 2007. We value teaching excellence, a solid research record, and professional experience. Contact: Dr. Richard Powell, Graziadio School of Business, Pepperdine University, 24255 Pacific Coast Highway, Malibu, CA 90263; Phone: (310) 506-4874; Email: rpowell@pepperdine.edu. Pepperdine is religiously affiliated with the Churches of Christ. It is the purpose of Pepperdine to pursue the very highest employment and academic standards within a context which celebrates and extends the spiritual and ethical ideals of the Christian faith. While students, faculty, staff, and members of the Board of Regents represent many religious backgrounds, Pepperdine reserves the right to seek, hire, and promote persons who support the goals and mission of the institution, including the right to prefer co-religionists. Pepperdine University is an Equal Opportunity Employer and does not unlawfully discriminate on the basis of any status or condition protected by applicable federal or state law.

536 SAN JOSE STATE UNIVERSITY, invites applications for Fall 2007. Four tenure-track accounting positions are available as Professor, Associate, Assistant, or temporary appointment. Positions are for faculty with an emphasis in Financial Accounting, Managerial Accounting, Audit/Systems, and/or Taxation. All candidates must have demonstrated versatility and competencies in subject matter and pedagogy. The teaching assignments are at both graduate and undergraduate levels, including traditional and accelerated formats. All positions are subject to final approval by the University. The minimum credential requirement for tenure-track appointments is a Ph.D or a Ph.D. candidate. Strong preference will be given to candidates who hold a Ph.D. from an AACSB-accredited university. The application deadline is December 31,2006 or until the position is filled. Send application letter, *vita*, teaching evaluations, and three letters of reference to: Dr. Janis Zaima, Accounting and Finance (BT850), San Jose State University, One Washington Square BT 850, San Jose, CA 95192-066. San Jose State University is an Equal Opportunity/Affirmative Action Employer. Interviews may be scheduled for AAA with Dr. Janis Zaima at (408) 924-3490 or zaima_j@cob.sjsu.edu.

537 WESTERN KENTUCKY UNIVERSITY, Ford College of Business invites applications for a tenure-track position in auditing, with secondary specialization open, beginning Fall 2007. Candidates with appropriate credentials may be considered for the Meany Professorship, which includes a salary stipend and an annual expense allowance, and may include tenure. Ph.D. required (ABD considered) and certification desirable. Applicants should have a proven record of quality teaching and research. Qualifications for the endowed professorship include a documented record of effective teaching and a distinguished record of research. Endowed professors are expected to provide leadership in scholarly activity and collaborate with colleagues on research. Applicants should submit a letter of application, *vitae*, transcripts, and three letters of reference to: Accounting Search Committee, Grise Hall 501, 1906 College Heights Boulevard, Bowling Green, KY 42101. Please indicate in your application letter if you wish to be considered for the Meany Professorship. Materials should be submitted by October 16, 2006 and the position will remain open until filled. Western is an Equal Opportunity Employer. Website: http://www.wku.edu/gfcb.

538 JACKSONVILLE UNIVERSITY, Davis College of Business invites applications for a tenure-track faculty position at the Assistant/Associate/Full Professor rank, effective Fall 2007. Preference will be given to those with related interests in Finance, Ethics, and/or Entrepreneurship. Candidates for the position must have a Ph.D. or a D.B.A. in Accounting or be significantly along in the dissertation process. Demonstrated excellence in teaching and the ability to foster productive partnerships on campus as well as in the business community are considered positive complements to the academic qualifications. The candidate will teach undergraduate and graduate courses in a five-year Accounting program designed to prepare students to sit for the CPA exam in Florida. Please send a letter of application, *vita*, transcripts, and contact information for three references to: DCOBFacultySearch@ju.edu; or to: Ms. Patty West, Jacksonville University, 2800 University Ave North, Jacksonville, FL 32211. Jacksonville University is an Equal Opportunity Educator and Employer.

539 WESTERN WASHINGTON UNIVERSITY, invites applications for tenure-track Assistant Professor positions in Auditing and/or Information Systems beginning Fall 2007. Doctoral degree in accounting required by appointment date. Duties include teaching undergraduate and graduate courses, university service, and scholarly publication. View announcement, including complete list of qualifications, at http://www.acadweb.wwu.edu/hr/Jobs/faculty.asp. Western overlooks Puget Sound, is situated between Seattle and Vancouver, B.C., and has 12,000 students. AACSB-accredited. Please send letter of application, *curriculum vitae*, transcripts, evidence of teaching effectiveness, and names of three references to: Professor Julie A. Lockhart, Chair, Department of Accounting, Western Washington University, Bellingham, WA 98225-9071; Phone: (360) 650-3202; Email: Julie.Lockhart@wwu.edu. Application Priority Review: 10/16/06. AA/EOE.

540 LEHIGH UNIVERSITY, seeks a scholar interested in joint tax research with other areas, including financial accounting, finance, and economics, and who also has a commitment to teaching taxes, to join its Accounting faculty in Fall 2007. The opening is for a graduating Ph.D. or senior Assistant Professor. A college, ranked the 18th best undergraduate B-School by *BusinessWeek* and with a recently acquired Perella endowment of $10M, a very successful M.S. in Accounting program, and a dynamic Accounting Department make this a great time to consider Lehigh. Applications will be accepted until the position is filled. If interested, send your *vita*, names of three references, and examples of scholarly work to Dr. Kenneth P. Sinclair, Chair, Department of Accounting, Lehigh University, 621 Taylor Street, Bethlehem, PA 18015-3117; Email: kps1@lehigh.edu. Qualified women and minority candidates should apply. Lehigh University is an Equal Employment/Affirmative Action Employer.

541 UNIVERSITY OF MAINE, Business School invites applications for two tenure-track positions at the Assistant Professor rank in Accounting beginning Fall 2007. Applicants must have or anticipate completion of a Ph.D. in accounting by Fall 2007. Teaching: (1) Audit and one other area of accounting, and (2) Financial and one other area of accounting. A balance between teaching and research is emphasized. Maine is a beautiful state with a national park and many recreational opportunities. For further details see: http://www.umaine.edu/eo/faculty/. Send a letter of application, *curriculum vitae*, three reference letters and student evaluations to: Professor Steven Colburn, Chair, MBS, University of Maine, 5723 DP Corbett Business Building, Orono, ME 04469-5723; Phone: (207) 581-1982; Email: Colburn@maine.edu. Review of applications will begin immediately and continue until the position is filled. UM is an EO/AA employer.

542 UNIVERSITY OF SOUTH FLORIDA, School of Accountancy invites applicants for two tenure-earning Assistant/Associate Professors beginning Fall 2007. These positions will involve teaching in undergraduate and master's programs and involvement with Ph.D. program. Other duties include conducting research leading to publications in refereed journals and service to the academic and professional communities. Ph.D. or near completion and specialization in financial accounting, managerial accounting or AIS is required. To apply or for more complete description of position including minimum and preferred qualifications contact: Robert M. Keith; School of Accountancy; College of Business Administration; University of South Florida; 4202 E. Fowler Avenue (BSN3403); Tampa, Florida 33620-5500; Phone: (813) 974-6516; Fax (813) 974-6528; Email: rkeith@coba.usf.edu. Applications close January 11, 2007. Position is subject to final university funding. USF is an Equal Opportunity/Affirmative Action Institution. For disability accommodations, call (813) 974-6389 a minimum of five working days in advance. Under Florida law, applications and meetings are public.

543 AMERICAN UNIVERSITY, Kogod School of Business invites applications for a Fall 2007 tenured or tenure-track appointment at the rank of Associate Professor or Professor, subject to final budgetary approval. Teaching interests should focus on managerial/cost accounting. Associate-level applicants must have strong research and teaching records. Professor-level applicants must have achieved a scholarly record of continuing distinction. Responsibilities include undergraduate and master's-level teaching, engagement in scholarly research, and participation in department and school activities. Salary and benefits are competitive. Electronic submissions preferred. Please send *curriculum vita,* three letters of recommendation, teaching evaluations, and recent published papers or working papers to: Jessica O'Connor, Kogod School of Business, 4400 Massachusetts Ave. NW, American University, Washington, D.C. 20016-8044; Email: oconnor@american.edu. For additional information, contact Robert Thompson at: rthomps@american.edu. Learn about the Kogod School of Business at: http://www.kogod.american.edu/. American University, an Equal Opportunity/Affirmative Action Employer, encourages applications from women and minorities.

544 SOUTHERN ILLINOIS UNIVERSITY CARBONDALE, School of Accountancy is accepting applications for a tenure-track position (Assistant or Associate Professor) starting January 1, 2007. Accounting specialty desired is Financial, but quality applicants in other areas will be considered. Applicant must have a strong commitment to a quality research program and excellent teaching. Applicants for senior rank must have an established research record in high-quality publications. Research responsibilities will include working with doctoral students. Teaching responsibilities will include participation in the undergraduate, M.Acc., and Ph.D. programs. Applicant should have completed all requirements for the doctorate or appropriate terminal degree at effective date of employment. Submit letter of application and resume to: Dr. Julie Sobery, Search Committee Chair, School of Accountancy, Mail Code 4631, Southern Illinois University Carbondale, 1025 Lincoln Drive, Carbondale, IL 62901; Phone: (618) 453-2289; Fax (618) 453-1411. Application deadline: September 30, 2006, or until position is filled. Southern Illinois University Carbondale is an Affirmative Action/Equal Opportunity employer that strives to enhance its ability to develop a diverse faculty and staff and to increase its potential to serve a diverse student population. All applications are welcomed and encouraged and will receive consideration.

545 UTICA COLLEGE, invites applicants for a tenure-track position in Financial Accounting at all ranks for Fall 2007. A research record commensurate with advanced rank is required for appointment at advanced rank. The primary area of specialization is Financial Accounting. Candidates should hold an earned doctorate from an AACSB-accredited institution. Professional certification is also desired. Candidates should demonstrate a strong interest in teaching and previous graduate teaching experience and distance learning experience is desired. Continued appointment requires evidence of scholarship and publication. Utica is the business and cultural center of the Mohawk Valley of upstate New York and is located in the foothills of the Adirondack Mountains. The area has excellent public schools, affordable housing, and ready access to all major northeastern cities by air, rail, and auto. Opportunities for both summer and winter recreation are abundant. Send *curriculum vita* and list of three current references by December 1 to: Dr. Hartwell C. Herring III, Search Committee Chair, Utica College, 1600 Burrstone Road, Utica, NY 13502. Applications from women and minorities are especially encouraged. The position will remain open until filled. AA/EOE.

546 TEXAS A&M UNIVERSITY, Department of Accounting in the Mays Business School invites applications for two new positions beginning in Fall 2007. The positions are open to rank and area of specialization, but we have a strong preference for applicants at either the advanced Assistant or Associate Professor level. Applicants should have a doctoral degree in accounting and should have a strong commitment to scholarly research and teaching excellence. Please send a cover letter and resume that includes: (1) educational achievement; (2) research/publication record; (3) indicators of teaching activity/effectiveness or potential; (4) work and other experience; and (5) a list of three references. All applications should be directed to: James Benjamin, Head, Department of Accounting, 4353 TAMU, College Station, TX 77843-4353; or Email: j-benjamin@tamu.edu). Review of applicants will begin on September 1, 2006 and will continue until the positions are filled. Texas A&M University is an Equal Opportunity/Affirmative Action Employer.

547 CALIFORNIA STATE UNIVERSITY, LONG BEACH, invites applicants for a tenure-track accounting faculty position, Assistant/Associate Professor for Fall 2007. Applicants must hold a Ph.D. or a D.B.A. degree with a major in Accounting and have demonstrated potential for excellence in teaching, research, curriculum development, and service contributions. Professional certification is desirable. Compensation will be competitive and be commensurate with the candidate's qualifications and experience. Applicant screening will begin November 6, 2006 and will continue until the position is filled. Please send a letter of application and *curriculum vitae* to: Dr. Steven A. Fisher Chair, Department of Accountancy, California State University, Long Beach, 1250 Bellflower Boulevard, Long Beach, CA 90840-8504; Email: sfisher@csulb.edu. CSULB is an urban university in the California State University system. It is located near the ocean in the City of Long Beach. CSULB is an Affirmative Action/Equal Opportunity Employer and has a strong commitment to the principle of diversity.

548 RUTGERS UNIVERSITY, CAMDEN, School of Business, invites applications for two tenure-track positions as assistant or associate professor in managerial/cost, financial/tax accounting or auditing. The candidates for assistant professor should have Ph.D. in accounting or should be near completion of their degree. Potential for quality research and effective teaching are expected. Candidates for associate professor must have Ph.D. in accounting and must have strong research and effective teaching record. Teaching load is 2/2. A letter of application and *curriculum vitae* should be submitted in writing and electronically to: Professor Kenis, Accounting Area Coordinator, Rutgers University, Camden Campus, School of Business, Camden., NJ 08102; Email: kenis@camden.rutgers.edu. The Camden Campus is located in southern New Jersey with easy access to Philadelphia, Atlantic City, and New York. Rutgers University is an Equal Opportunity Employer.

549 NATIONWIDE SEMINAR COMPANY, seeks experienced public speakers/CPA to lead the following topics: SEC filing and compliance; SOX updates; SOX Sec. 404 compliance; Fraud; Internal Control Design and Documentation; Using Excel for Accounting and Finance applications; A&A for industry; and Excel 2007. Must be able to lead a two-day seminar twice per month in major cities throughout the U.S. Top dollar for top talent. Respond to: careers@nccetraining.com.

550 UNIVERSITY OF NORTH CAROLINA, WILMINGTON, Department of Accounting invites applications for a tenure-track position at the Assistant or Associate level beginning Fall 2007. Background in Financial Accounting and previous related teaching and academic work experience preferred. Additional qualifications in Auditing or Not-for-Profit/Governmental desirable. Ph.D. in Accounting and a record of commitment to quality teaching and research required (ABDs considered if completed by 8/1/07). Apply online at" http://consensus.uncw.edu. For questions regarding the online application process, contact Gloria Sasser at (910) 962-3509. In addition to the online application requirements, three letters of reference must be received to complete the formal application process. Directions for submitting letters of reference will be emailed to applicants. Requests for an interview at the AAA meeting should be directed to Dan Ivancevich, at ivancevichd@uncw.edu by 7/25/06. Review of applications will begin on 11/1/06. UNCW is AACSB-accredited and an EEO/AA employer. Women/minorities are encouraged to apply.

551 DEPAUL UNIVERSITY, invites applications for the position of Director effective July 1, 2007. The ideal candidate will possess a strong record of scholarship, teaching, and administrative skills, including effective administration at the department chair or dean level or an administrative position that includes performance appraisal of faculty. The preferred candidate should have an accounting or MIS doctorate from an AACSB-accredited institution, relevant professional certification(s), professional experience in accounting or MIS, personal capabilities with computer literacy, effective communication skills, excellent organization skills and leadership skills, knowledge and experience with AACSB-accreditation standards, strong relationship building and fundraising skills, and the ability to promote, develop, and create an atmosphere for quality teaching, research, and service. Candidates should email an application letter, *vita*, and at least three references (preferably using Microsoft Word) to: John Ahern, Chair, Director Search, Email: jahern@depaul.edu. For more information, visit: http://accountancy.depaul.edu.

552 UNIVERSITY OF MIAMI, Department of Accounting is seeking applicants at the Assistant Professor level for a tenure-track position beginning in the Fall 2007. The individual will be expected to teach at both the undergraduate and graduate levels, and to conduct and publish significant research in their subject of interest, focusing primarily on important journals in the accounting literature. While applications in all areas will be welcome, our area of most pressing need is auditing. Candidates should currently expect to have completed the Ph.D. or equivalent by the time of their appointment as University of Miami faculty. Relevant professional certification is desirable. Interested candidates should contact Dr. Oscar Holzmann, Chair of the Faculty Recruiting Committee either by mail or email: Dr. Oscar Holzmann, Department of Accounting, 309 KE Building, University of Miami, Coral Gables, FL 33146-6531; Email: holzmann@miami.edu.

553 WASHINGTON STATE UNIVERSITY, VANCOUVER, invites applications for an Assistant Professor of Accounting faculty position, beginning August 16, 2007. A doctorate or ABD in Accounting by the first day of employment, and demonstrated record of published research or research promise is required. Successful candidate will be expected to teach undergraduate tax accounting as well as introductory accounting, and a strong global perspective is desirable. Successful candidate will receive a six-hour semester teaching load, summer research support, competitive salary, and excellent fringe benefits. Review of applications begins August 15, 2006. Electronically submitted application materials are preferred: acctg@vancouver.wsu.edu; Accounting Search, Business Programs, Washington State University, 14204 NE Salmon Creek Avenue, Vancouver, WA 98686. Further information regarding application, position and contact: http://www.vancouver.wsu.edu/programs/bus/acctgsearch.html. WSU is an EEO employer. Protected groups are encouraged to apply.

554 BUTLER UNIVERSITY, invites applications for a tenure-track position at the Assistant or Associate rank, starting August 2007, subject to final budget approval. Candidates should have research and teaching interests in either Managerial or Financial Accounting. Applicants should possess a Ph.D., be capable of high quality teaching, and have an active research program. The college is AACSB-accredited with 625 undergraduates and 200 graduate students. We recently received a $22 million grant from the Lilly Endowment to make our programs more relevant, integrated and experiential. This grant provides opportunities for program development, faculty development, research support, and interaction with the business community. Butler is a selective, comprehensive, private university of 3,800 students set on a wooded 300 acre campus just north of Indianapolis. Candidates should send a current *vita*, along with a letter of intent to: Russ Kershaw, Associate Dean of Academic Programs, College of Business Administration, Butler University, 4600 Sunset Avenue, Indianapolis, IN 46208; Email: rkershaw@butler.edu; Phone: (317) 940-9841. Butler University is an Equal Opportunity Employer and is committed to enhancing the diversity of the student body and the faculty and staff. Women and minorities are strongly encouraged to apply.

555 MORGAN STATE UNIVERSITY, Graves School of Business and Management invites applications for one tenure-track position at the Associate Professor or Assistant Professor rank starting August 2007 contingent on budget approval. Candidates should have a Ph.D. degree in accounting from an AACSB-accredited institution and research and teaching interests in financial accounting or managerial accounting. Candidates for Associate Professor should have a significant record of publications in recognized accounting journals. Salaries are AACSB competitive. The University has business school and accounting accreditation by the AACSB. The business school offers a B.S. degree in accounting and a Ph.D. degree in Business Administration/Accounting. Master's level accounting courses are offered in the general M.B.A. program. Interested candidates should send academic *vitae* to: Alex P. Tang, Chair, Recruiting Committee, Department of Accounting and Finance, Morgan State University, Baltimore, MD 21251; or Email: atang@moac.morgan.edu. Morgan State University is an Affirmative Action/Equal Opportunity Employer.

556 MISSISSIPPI STATE UNIVERSITY, College of Business and Industry, School of Accountancy, invites applicants to apply for a tenure-track position with a specialization in financial accounting to begin Fall 2007. Rank is open, but candidates must have an accounting doctorate degree, demonstrated excellence in teaching, publication record in nationally refereed journals, and experience in working with doctoral students. A primary responsibility is to assist with leadership/involvement with the School's doctoral program. In addition, the candidate is expected to participate in professional accounting activities and advise students. Professional certification and accounting experience are highly desirable. Salary is competitive and a named professorship may be available for candidates with outstanding credentials. Send letter of interest, current *curriculum vitae*, and names (including contact information) of three references to: Chair, Search Committee, School of Accountancy, Drawer EF, Mississippi State, MS 39762. Position will remain open until filled. MSU is an AA/EEO Employer.

557 GEORGIA INSTITUTE OF TECHNOLOGY, College of Management, is seeking candidates for the following positions starting August 2007: (1) one tenure-track faculty position. Advanced Assistant or Associate Professor candidates are preferred. A Ph.D. in Accounting or related discipline is required. Salary and teaching loads are very competitive; (2) one non-tenure-track Academic Professional position. The College offers academic programs leading to Bachelor's, M.B.A., and Ph.D. degrees. In addition, the College offers an Executive Master's degree in Management of Technology. For more information, please see: http://mgt.gatech.edu. Please send a letter of application, *curriculum vitae*, three reference letters, and for the tenure-track position, send copies of research papers to: Professor Arnold Schneider, College of Management, Georgia Institute of Technology, 800 W. Peachtree Street, Atlanta, GA 30308-0520. Applications will be reviewed beginning October 1, 2006. The deadline for submission is January 15, 2007. Georgia Tech is an Equal Opportunity/Affirmative Action Employer, and values diversity.

558 CALIFORNIA STATE UNIVERSITY, BAKERSFIELD, seeks candidates for tenure-track position in Accounting. Rank depends upon experience. Should possess earned doctorate from AACSB-accredited institution or anticipate completion by August 31, 2007. Ability to demonstrate teaching excellence is a must. Should be able to build or continue a research record consistent with the mission of this school and the requirements for maintaining its AACSB-accreditation. Teaching at both lower- and upper-division levels required. Demonstrated ability to teach at the M.B.A. level and professional certification (CPA, CMA, CIA) desired. Interested candidates should submit letter of application, complete *vita*, and three letters of reference: Dr. Janis Ruiz, BPA, California State University, Bakersfield 9001 Stockdale Highway, 20 BDC, Bakersfield, CA 93311-1099. Reviews of applications begin October 30, 2006. California State University, Bakersfield fosters and appreciates ethnic and cultural diversity among its faculty and students. Applications from women, ethnic minorities, veterans, and individuals with disabilities welcome.

559 UNIVERSITY OF DENVER, Daniels College of Business, School of Accountancy invites applications for a tenure-track position, rank open. Applicants should have a doctoral degree in accountancy and should have a primary teaching and research interest in financial accounting with a secondary interest in auditing desirable. The College encourages applications from those with cross-disciplinary interests, in the areas of ethics, internationalization, and innovation. The search begins immediately. The preferred starting date is September 1, 2007. Candidates must have a demonstrated record of excellent teaching and publication. The Daniels College of Business and the School of Accountancy are AACSB-accredited. To apply, see: http://www.du.edu/hr/employment/jobs.html. For further information, contact: Dr. John Bazley, School of Accountancy, University of Denver, Denver, CO 80208; Phone: (303) 871-2023; Email: jbazley@du.edu. The University of Denver is an EEO/AA employer and is committed to the diversity of its faculty and staff and encourages applications from women, minorities, people with disabilities and veterans.

560 TARLETON STATE UNIVERSITY, invites applications for a full-time, tenure-track position in accounting or related area. Earned doctorate required in accounting or in a closely related field. Position available January 2007. Send letter of interest, *vitae*, and copy of doctoral transcript to: Dr. John P. Idoux, Tarleton-Central Texas, 1901 South Clear Creek Road, Killeen, TX 76549. Visit T-CT at: http://www.tarleton.edu/centraltexas. Tarleton, a member of The Texas A&M University System, is an Equal Employment Opportunity/Affirmative Action Employer.

561 SHIPPENSBURG UNIVERSITY, is seeking candidates for the position(s) of Assistant Professor of Accounting, effective January or August 2007. Associate Professor level may be considered, depending on qualifications. Refer to the University website to view the full ad for requirements and qualifications: http://www.ship.edu/HR/index.html. Qualified candidates should send a written letter of application, current *curriculum vitae*, official undergraduate and graduate transcripts, and three written letters of reference to Dr. Scott N. Cairns, John L. Grove College of Business, Shippensburg University, 1871 Old Main Drive, Shippensburg, PA 17247-2299. Applications will be considered complete when all of the above have been received. Applications are reviewed upon receipt as part of an ongoing search and will continue to be reviewed until the position(s) is filled. All applicants must furnish proof of eligibility to work in the U.S. upon appointment. Shippensburg is an Equal Opportunity Employer.

562 UNIVERSITY OF NEW HAMPSHIRE, Whittemore School of Business and Economics invites applications for one (possibly two) tenure-track positions in accounting at the Assistant Professor level beginning Fall 2007. Applicants must have a doctoral degree or be near completion. Teaching interests in managerial accounting and/or accounting information systems are preferred, however candidates with interests in other areas will be considered. Faculty members are expected to publish in peer-reviewed journals. Salary and fringe benefits are competitive and the school subscribes to various financial databases through WRDS. Interested applicants should mail a packet containing a letter of interest, current *vita*, three recommendation letters, recent teaching evaluations, and evidence of scholarly work to: Professor Afshad Irani, Box A, McConnell Hall, University of New Hampshire, Durham, NH 03824; Phone: (603) 862-3342. The School is accredited by the AACSB and is an Affirmative Action/Equal Opportunity Employer with a commitment to excellence through diversity.

563 UNIVERSITY OF ILLINOIS AT URBANA–CHAMPAIGN, Department of Accountancy is recruiting faculty for full-time appointments at all ranks. Candidates at advanced ranks must possess a doctoral degree, and candidates for assistant professor positions must possess or be near completion of a doctoral degree. Applicants for senior positions should have a strong research record and the ability to participate in doctoral education and curriculum development at the undergraduate and graduate levels. In addition, non-tenure-track Lecturer and visiting positions are available. Applicants should submit a resume that lists research publications and courses taught. Compensation is competitive and will be commensurate with the appointment level and applicant's record. To ensure full consideration, applications must be received by November 1, 2006. Applications received thereafter will receive full consideration if there are positions still available. Proposed starting date is May 16, 2007 or as soon as possible thereafter. Correspondence and resumes should be sent to: Recruiting Chair, Department of Accountancy, 360 Wohlers Hall, College of Business, University of Illinois, 1206 S. Sixth Street, Champaign, IL 61820. The University of Illinois is an Affirmative Action/Equal Opportunity Employer.

564 JAMES MADISON UNIVERSITY, School of Accounting invites applications for an Assistant or Associate Professor tenure-track position starting August 2007. Candidates should have a teaching interest in Accounting Information Systems, Financial, or Managerial, doctorate in Accounting from an accredited school, and prospects for publication and excellence in teaching. The School offers undergraduate and master's degrees in accounting and is an academic unit of the College of Business. The College of Business undergraduate program is ranked 35th by *BusinessWeek.* JMU has an online application for faculty positions. To begin the process, paste the following into the address field of your browser: http://joblink.jmu.edu. The position number is 0400831. Address questions to: Paul Copley, Director; Phone (540) 568-3081; Email: copleypa@jmu.edu. JMU is an Equal Opportunity/Affirmative Action/Equal Access Employer and especially encourages applications from women, minorities, and persons with disabilities. All offers of employment are contingent up on a criminal history check.

565 ROCHESTER INSTITUTE OF TECHNOLOGY, seeks applications for an Assistant or Associate Professor of accounting starting September 2007. Research and graduate/undergraduate teaching in management accounting/cost management and one other area (preferably AIS or auditing or tax) to meet program needs. Candidates must hold Ph.D. in accounting from an AACSB-accredited institution. ABDs with expected completion by January 2008 considered. Associate Professor requires established record of teaching and research. Professional experience and/or professional certification, and a willingness to interact with the professional community are desirable. Candidates must also have strong research and interpersonal skills and an ability to function as a team member. The College of Business, with 45 faculty, offers B.S., M.S., and M.B.A. degrees. RIT, a private university with 15,000 students from 100 countries, provides technology-based educational programs for personal and professional development. Send inquiries for nominations and applications (with *vita* and reference information) to: Professor Khondkar Karim, Chair, Accounting Faculty Search Committee, E. Philip Saunders College of Business, Rochester Institute of Technology, Rochester, NY 14623-5608. The position will remain open until filled. RIT is an Equal Opportunity/Affirmative Action Employer. Email: kkarim@cob.rit.edu. Website: http://www.cob rit.edu.

566 WASHINGTON STATE UNIVERSITY, invites applications for a full-time, permanent, tenure-track accounting faculty position at the Assistant or Associate level, beginning August 16, 2007, at the Pullman, WA campus. A doctorate or ABD in Accounting or related field before date of hire, a record of published research or research potential, and an interest in teaching auditing and/or financial accounting is required. The ability to teach auditing and/or financial accounting is desirable, as is a strong global perspective. The successful candidate will receive a six-hour semester teaching load, summer research support, competitive salary, and excellent fringe benefits. Review of applications begins September 15, 2006. Send letter of application, résumé, student teaching evaluations (or summary), recent research paper, and three letters of recommendation to: Professor Richard Toolson, Accounting Search Chair, Department of Accounting, Box 644729, Washington State University, Pullman, WA 99164-4729. WSU is an EEO employer. Protected groups are encouraged to apply.

567 BAYLOR UNIVERSITY, Department of Accounting and Business Law at invites applications for two tenure-track positions at all ranks for Fall 2007. Preference will be given to those applicants with teaching interests in auditing, financial reporting and taxation. A Ph.D. (or ABD) in accounting is required; professional certification is preferred. Applicants must demonstrate research and teaching excellence. Applications, including *curriculum vitae*, three letters of reference, working paper or defended dissertation proposal, and evidence of an active Christian faith commitment should be sent to Dr. Charles E. Davis, Department of Accounting and Business Law, Baylor University, One Bear Place #98002, Waco, TX 76798-8002; Email: Charles_Davis@baylor.edu. Applications will be reviewed beginning October 1, 2006 and will be accepted until the position is filled. Baylor University is a Baptist university affiliated with the Baptist General Convention of Texas. As an Affirmative Action/Equal Employment Opportunity Employer, Baylor University encourages minorities, women, veterans, and persons with disabilities to apply.

568 BLOOMSBURG UNIVERSITY, Department of Accounting, AA#91-5-243. One tenure-track position at the Assistant/Associate level beginning Fall 2007. The Department offers a Bachelor's degree in Accounting and a Concentration in Fraud Examination. AACSB International accredited. Earned doctorate preferred; ABD considered with a contingency contract. Consideration will also be given to those that meet AACSB standards for professional qualification. Prior accounting teaching experience, professional work experience, ability to meet College of Business standards for scholarly activity, and/or CPA, CMA, CIA, CFE certifications are desirable. Ability to work with diverse populations preferred. Review of applications will begin November 1, 2006. The position will remain open until filled. For a full position description, including application procedures, please visit: http://www.bloomu.edu/jobs. AA/EEO.

569 COLLEGE OF CHARLESTON, Charleston, South Carolina invites applications for Assistant or Associate Professor tenure-track openings for accounting faculty beginning spring or Fall 2007. Two positions are available; primary needs are in Financial Accounting, Information Systems, and Governmental/Not for Profit Accounting. The College of Charleston offers the BS in Accounting and Master of Science in Accountancy degrees. Information about our program can be found at: http://www.cofc.edu. Candidates are expected to have a Ph.D. degree from an AACSB-accredited university by the time employment begins. Teaching experience, business experience, and professional certification are desirable. Applicants should send a cover letter and *vitae* to: koprowskiw@cofc.edu or to: William Koprowski, Department of Accounting and Legal Studies, 66 George Street, Charleston, SC 29424; Phone: (843) 953-7835. The College of Charleston is an Affirmative Action/Equal Opportunity Employer.

570 STANFORD UNIVERSITY, Graduate School of Business, invites applications for tenured and non-tenured faculty positions in Accounting, effective September 2007, from candidates at all levels and stages of their careers who are interested in financial accounting, managerial accounting, tax, and other areas. Candidates must have a Ph.D. and the highest level of accomplishment or promise in both scholarly research and M.B.A. and Ph.D. teaching. Applications should include *curriculum vitae* and copies of research publications. Applicants are strongly encouraged to submit applications by uploading *curriculum vitae*, publications, and teaching evaluations (if applicable) to: http://www.gsb.stanford.edu/recruiting. Alternatively, you may email materials to: faculty_recruiter@gsb.stanford.edu; or mail non-returnable materials to: Faculty Recruiting Manager, Box A, Graduate School of Business, Stanford University, Stanford, CA 94305-5015 (please allow three weeks for processing). Three letters of recommendation should be sent to that same address. The deadline is January 19, 2007. Stanford University is an Equal Opportunity/Affirmative Action Employer.

571 UNIVERSITY OF CALIFORNIA, SAN DIEGO, The Rady School of Management (http://management.ucsd.edu/) of the University of California, San Diego (UCSD) invites applications for Assistant, Associate and Full Professor positions in accounting. Preference will be given to experienced applicants who have distinguished scholarly records, have demonstrated teaching expertise with graduate students and executives, demonstrated leadership efforts and/or contributions to diversity and are good institution builders. All applicants must have a Ph.D. or will be working toward a Ph.D. by the start date of the new academic year. Associate-level candidates must show evidence of a strong research record in their specialization, while candidates for senior rank must demonstrate a continuing publication record in leading journals. Salary and appointment level are dependent on experience and based on University of California pay scales. The positions are expected to have a start date of July 1, 2007. A detailed vita (pdf or Word document form) will be accepted by email only to: radyaccounting@ucsd.edu. Please indicate reference Job Code: ARAACCT on the subject line and state the level of position you are applying for in your cover letter. Review will begin October 1, 2006 and will continue until positions are filled. UCSD is an Affirmative Action/Equal Opportunity Employer with a strong institutional commitment to the achievement of excellence through diversity among its faculty and staff.

572 UNIVERSITY OF NORTH ALABAMA, invites applications for Raburn Eminent Scholar of Accounting at the rank of Professor. Requirements include earned Ph.D. and professional certification in accounting, active continuing research agenda, and commitment to teaching. Candidates from all accounting disciplines may apply, but auditing candidates are preferable. Professional accounting experience is highly desirable. The Raburn Chair holder will provide leadership in teaching, research, and service; promote Accounting Department and College of Business in local, regional, and national forums; and help facilitate AACSB-accreditation for the College. Application review begins immediately and continues until position is filled in 2007. Submit application letter, detailed *vitae*, unofficial copies of graduate and undergraduate transcripts, and complete contact information for three references to: Director of Human Resources and Affirmative Action, University of North Alabama, UNA Box 5043, Florence, AL 35632-0001. UNA is an Equal Opportunity Employer committed to achieving excellence and strength through diversity.

573 SAN DIEGO STATE UNIVERSITY, School of Accountancy invites applications for a tenure-track position in audit/systems, financial, tax, or international at the assistant/associate level. Applicants should complete a Ph.D. or D.B.A. by August 2007. Appointments at advanced rank require commensurate experience. Preference is given to candidates with a multidisciplinary and integrative approach to accounting. Professional certification and experience are desirable. Positions are contingent upon state funding. Detailed descriptions are available at: http://www.sdsu.edu/accounting. If interested, email or send a letter with your current *curriculum vitae*, three recommendation letters, and evidence of teaching effectiveness and scholarly contributions to: Sharon Lightner, Director School of Accountancy, SDSU San Diego, CA 92182-8221; Phone: (619) 594-3735; Fax: (619) 594-3675; Email: sharon.lightner@sdsu.edu. SDSU is a Title IX, Equal Opportunity Employer and does not discriminate against individuals on the basis of race, religion, national origin, sexual orientation, gender, marital status, age, disability, or veteran status, including veterans of the Vietnam era.

574 GRAND VALLEY STATE UNIVERSITY, Department of Accounting and Taxation in the Seidman College of Business invites applications for a new position beginning in Fall 2007. The position is open to rank and area of specialization. Applicants should have a doctoral degree in accounting and a strong commitment to teaching excellence and scholarly research. The Department and College are both accredited by the AACSB. Salary is competitive. Please send a cover letter and resume that includes: (1) educational achievement; (2) research/publication record; (3) indicators of teaching effectiveness or potential; (4) work and other experience; and (5) a list of three references to: Ms. Kathy Goralski, Accounting Search Committee, Department of Accounting and Taxation, Grand Valley State University, 401 West Fulton Street, Grand Rapids, Michigan 49504; or Email: goralskk@gvsu.edu. Review of applicants will begin on September 1, 2006 and will continue until the position is filled. Grand Valley State University is an EO/AA institution.

575 LAM RESEARCH, in Fremont, California invites candidates to apply for our Director, Tax opportunity, job ID#1621. The candidate's responsibilities will include the development and management of corporate strategic tax vision objectives; development of corporate/organizational policies and authorizes/oversees their implementation; tax planning, ensuring timely tax compliance filings and execution of tax strategies; development and implementation of systems to compile information for various federal, state, and local reports; minimizing corporate tax through use of tax laws and regulations, organization's compliance with all international, federal, state, local, and applicable foreign tax laws and regulations; and tax research programs to ascertain the minimum tax liability for the organization. Accounting and CPA education is required. Qualifications also include external reporting tax knowledge and experience, internal audit experience, and at least 15 years' experience working with and supervising staff in a Big 4 accounting firm or large multinational company. Experience with FAS No. 109 required. Please send resumes directly to: kelly.orasin@lamrc.com, http://www.lamrc.com.

576 UNIVERSITY AT BUFFALO, School of Management seeks a tenured senior research faculty member in accounting, beginning August 2007. Teaching and research interests should be in financial accounting, managerial accounting, or auditing. Candidate will be expected to take a leadership role in mentoring our financial/capital markets-based junior faculty and doctoral students. Candidates must have a substantial record of scholarly research in leading accounting journals and quality teaching at the undergraduate, M.B.A. and/or Ph.D. level. Salary and fringe benefits are highly competitive. Teaching opportunities are available at the Ph.D., M.B.A., Executive M.B.A., International M.B.A., M.S. in accounting, and undergraduate degree programs. Send *vita* and supporting information to: Professor Ronald Huefner, Recruiting Chair, Department of Accounting and Law, 370 Jacobs Center, University at Buffalo, Buffalo, NY 14260-4000; or Email: rhuefner@buffalo.edu. The University at Buffalo is an Affirmative Action/Equal Opportunity Employer.

577 UNIVERSITY AT ALBANY, Department of Accounting invites applications for multiple tenure-track positions beginning at the Assistant Professor level starting September 2007. The candidates should have a doctoral degree in accounting or be ABD with expected completion by December 2007. Candidates with doctoral degrees in allied disciplines such as Economics, Computer Science, Statistics, and Criminal Justice are also strongly encouraged to apply provided they have strong research interests in accounting, assurance, and/or forensics related areas. Evidence of a strong commitment to scholarly research and excellence in teaching, and a demonstrated ability to work with and/or instruct diverse groups of people is required. Candidates in all areas of accounting will be considered. Candidates are requested to submit their *vitae* and contacts for three references by email to: jgangolly@uamail.albany.edu; Jagdish S. Gangolly, Accounting Search Chair, Department of Accounting and Law, School of Business, University at Albany, Albany, NY 12222. The job announcement is available at: http://hr.albany.edu/vacancy/faculty/P06-65.html.

578 RICE UNIVERSITY, Jones Graduate School of Management, invites applications for a tenure-track position in Financial Accounting, effective Fall 2007. The position is open rank, but applicants at the Associate or Full level with demonstrated excellence in scholarly research and teaching are especially encouraged to apply. Applicants at the Assistant level must have a Ph.D. or be near completion, and should demonstrate strong potential for quality research and teaching. Interested candidates should submit a cover letter, *curriculum vitae*, and copies of recent research and teaching evaluations to: accountingrecruit@rice.edu; or to: Ms. Annepha Pemberton, Accounting Recruiting, Jones Graduate School of Management, Rice University, PO Box 2932, Houston, TX 77252-2932. Review of applications will begin October 1, 2006 and continue until the position is filled. If you have further questions, please contact Karen Nelson at: nelsonk@rice.edu. Rice University is an Equal Opportunity/Affirmative Action Employer.

579 UNIVERSITY OF UTAH, School of Accounting and Information Systems invites applications for new tenure-track faculty positions in Accounting at all ranks in any area of specialization beginning Fall 2007; pending funding and University approval. Applicants should possess a doctorate or be close to completion. We offer competitive salaries, research support, teaching loads, and excellent fringe benefits. The School of AIS and the David Eccles School of Business are accredited by the AACSB. Send application by January 15, 2007, including dissertation proposal or working paper, *vita*, and three letters of reference to: Robert Allen, Director, School of AIS, 1645 E. Campus Center Dr., Room 108, University of Utah, Salt Lake City, UT 84112. Late applications accepted until position is filled. The University of Utah is an Equal Opportunity/Affirmative Action Employer, encourages applications from women and minorities, and provides reasonable accommodation for the known disabilities of applicants/employees.

580 SABANCI UNIVERSITY, Faculty of Management (FMAN) invites applications for full-time and visiting positions in all areas of Accounting and Finance for the 2007–2008 academic year. Sabanci University, a private, innovative academic institution, offers excellent facilities and a light teaching load amid a modern campus located in Istanbul, one of the world's truly cosmopolitan cities. The university provides faculty members excellent support, including housing on campus or a housing stipend. FMAN offers undergraduate, M.B.A., executive M.B.A., and doctoral degree programs, with English as the medium of instruction. For further information, please visit: http://www.sabanciuniv.edu/ybf/eng/ Applicants are asked to complete an online application form at: http://www.sabanciuniv.edu/English/?BilgiHizmetleri/Ik/Ik.html. Please include ACC006 as the Reference Code. Additionally, *curriculum vitae*, a statement of research and teaching activities, and the names of three references should be sent via email to: nakiye@sabanciuniv.edu (Nakiye Boyacigiller, Dean).

581 YONSEI UNIVERSITY, School of Business, Seoul, Korea, invites applications for Spring or Fall 2007, tenure-track faculty positions in Accounting at all levels, subject to final budgetary approval. Both Korean and foreign candidates are encouraged to apply for the positions. Teaching interests should focus on managerial accounting, financial accounting or auditing. Candidates must have a Ph.D (or D.B.A.) or be near completion of the dissertation with a concentration in any area of accounting. Compensation will be highly competitive, commensurate with qualifications and experience. To be considered, please send *curriculum vita*, three letters of recommendation, teaching evaluations, and recent published papers or working papers to: Dr. Jinho Chang, 134 Seodaemoon-Ku, Shinchon-Dong, School of Business, Yonsei University, Seoul, Korea, 120-749. For additional information, contact Dr. Jinho Chang at: chang@yonsei.ac.kr

582 LOUISIANA STATE UNIVERSITY, E. J. Ourso College of Business seeks to fill the Ourso Distinguished Chair in Accounting at the rank of Associate or Full Professor. This endowed chair was established by Mr. E. J. Ourso and his wife, Marjory B. Ourso, to support an eminent scholar in the field of accounting. Required Qualifications: earned doctorate in accounting; established record of high-quality academic research; continuing research agenda; interest in guiding the future of the Ph.D. program in accounting. Applications and nominations will be accepted through November 10, 2006, or until candidate is selected, which could be as early as August 21, 2007. An offer of employment is contingent on a satisfactory pre-employment background check. Letter of application, *curriculum vitae* (including email address), and/or nominations should be sent to: Professor Larry Crumbley, Chair of the Search Committee, Department of Accounting, E. J. Ourso College of Business, CEBA 3106, Louisiana State University, Ref: Log #0690, Baton Rouge, LA 70803; Email: dcrumbl@lsu.edu. Information about the Department of Accounting can be found at: http://www.bus.lsu.edu/academics/accounting/. LSU is an Equal Opportunity/Equal Access Employer.

583 THE GEORGE WASHINGTON UNIVERSITY, Department of Accountancy, School of Business invites applications for appointment as a tenure-track Assistant Professor starting fall 2007. Applicants should have completed a Ph. D. degree in accountancy or a related field by August 1, 2007. ABDs will be considered at the rank of Instructor and, if appointed, must complete the doctoral degree by May 31, 2008. The successful candidate will have strong potential for scholarship as evidenced by current scholarly working paper(s) and/or publications, and teaching excellence as indicated by research presentations and teaching assessments. Salary and benefits are competitive. Application procedure: Send an application letter, resume, and a current working paper to: Professor Keith E. Smith, Accountancy Chair, School of Business, The George Washington University, Washington, D.C. 20052; or Email: accyjob@gwu.edu. Applications review begins October 2, 2006 and continues until the position is filled. The GWU is an EO/AAE.

584 VANDERBILT UNIVERSITY, Owen School of Management, Nashville, TN, invites applications at all levels for one tenure-track accounting faculty position beginning September 1, 2007. The primary need is for an individual with expertise in financial reporting; we will consider qualified individuals in auditing, tax, or managerial accounting. All candidates must have a Ph.D. and promise for high-level accomplishments in research and teaching. Significant research and teaching records are required for tenured positions. Rank and salary will be commensurate with qualifications. Please send application letter, *curriculum vita*, representative working papers and publications, teaching evaluations (if available), and names or letters of reference to: Professor Richard Willis, Vanderbilt University, The Owen School of Management, 401 21st Avenue South, Nashville, TN 37203. Vanderbilt University is an Affirmative Action/Equal Opportunity Institution.

585 WASHINGTON UNIVERSITY IN ST. LOUIS, Olin School of Business is seeking faculty in accounting at all levels beginning in August 2007. Candidates for Associate or Full Professor should have an established record of quality research, while candidates for Assistant Professor should be promising research scholars. Successful candidates will be expected to publish high-quality research and teach effectively in the school's Bachelor's, Master's, and/or Ph.D. programs. Candidates should hold a doctorate degree or be close to completion. Salary, teaching load, and faculty support are highly competitive. Applicants should send a cover letter, *curriculum vitae*, and a copy of the dissertation proposal or completed dissertation, or in the case of more established scholars, several samples of the applicant's scholarship. Applications should be sent to: Dean Mahendra R. Gupta, Olin School of Business, Campus Box 1133, Washington University, One Brookings Drive, St. Louis, MO 63130-4899. Early applications receive fullest consideration. Washington University is an Equal Opportunity/Affirmative Action Employer and welcomes applications from women and minorities.

586 BARUCH COLLEGE, Stan Ross Department of Accountancy, invites applications for anticipated openings for tenure-track and visiting positions at all ranks. Areas of specialization include financial and managerial accounting, taxation, auditing (internal and external), and information systems. Rank depends on qualifications and salaries are competitive. Candidates should have a doctoral degree in Accountancy or be close to the completion of that degree. For ranks above Assistant Professor, candidates should have a completed Ph.D. degree, a record of excellent scholarly research and demonstrated teaching effectiveness. Please submit *curriculum vitae*, a research paper, and letters of reference by February 15, 2007 to: Professor Donal Byard, Stan Ross Department of Accountancy, One Bernard Baruch Way, New York, NY 10010. Baruch College is an AA/EO/IRCA/ADA Employer.

Tenure Track Position

The School of Accounting at Florida Atlantic University invites applications to fill a tenure-track position at the Assistant/Associate level starting August 2007. Our greatest needs are in the areas of Financial Accounting, Auditing, AIS and Managerial Accounting—but candidates in other areas are also encouraged to apply. Candidates must have earned a Ph.D. or DBA or be close to completing the degree. Applicants for rank above Assistant Professor must have a distinguished record. The School offers degrees at the Bachelor, Master and Ph.D. levels and has executive education programs in Taxation and Forensic Accounting. The College of Business at FAU is accredited by the AACSB and salary is competitive. FAU faculty will be attending the National AAA Meeting in August 2006 and would like to schedule interviews with interested candidates. Review of applications will begin immediately and continue until the position is filled. Send application, curriculum vitae, and names of three references to Ms. Susan Dobson, Senior Secretary, School of Accounting, Florida Atlantic University, 777 Glades Road, Boca Raton, FL 33431. ***Electronic copies of the above are preferred*** and should be sent to dobson@fau.edu **and** sbhatt@fau.edu. FAU does not discriminate on the basis of race, color, national origin, sex, religion, age, disability, veteran status or sexual orientation in employment.

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I certify that all information furnished on this form is true and complete. I understand that anyone who furnishes false or misleading information on this form or who omits material or information requested on the form may be subject to criminal sanctions (including fines and imprisonment) and/or civil sanctions (including civil penalties).

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